VICTOR HENRIQUE LANA PINTO

ESSAYS ON COMMODITIES AND ECONOMIC UNCERTAINTY

Dissertation submitted to the Applied Economics graduate program of the Universidade Federal de Viçosa in partial fulfillment of the requirements for the degree of *Doctor Scientiae.*

Advisor: Leonardo Bornacki de Mattos

Co-advisor: Lorena Vieira Costa Lelis

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ABSTRACT

PINTO, Victor Henrique Lana, D.Sc., Universidade Federal de Viçosa, September 2024. **Essays on commodities and economic uncertainty.** Advisor: Leonardo Bornacki de Mattos. Co-advisor: Lorena Vieira Costa Lelis.

Commodities play a pivotal role in global trade, encompassing raw materials and agricultural products vital for economic growth and stability. This dissertation investigates the intricate relationship between commodities and economic uncertainty, exploring how these factors interact and influence economic dynamics. Drawing on a multidisciplinary framework, this study encompasses two independent papers that collectively contribute to the understanding of commodity markets and economic uncertainty. The first paper examines the impact of commodity exports on institutional quality across a range of developing and advanced economies from 1997 to 2022. Using panel data analysis and China's accession to the World Trade Organization as a natural experiment, the study reveals a significant negative causal relationship between commodity exports and institutional development. This finding underscores the adverse effect of commodity wealth on institutional effectiveness, highlighting the complexities of economic governance in resource-dependent nations. In the second paper, the focus shifts to the interactions between commodity futures (specifically live cattle and soybeans) and the stock market in Brazil from January 2014 to July 2023. Through time-series analysis, we investigate the influence of policy changes on commodity-stock dynamic correlations. We do not observe an economically significant effect of economic uncertainty originated in Brazil and abroad on the time-varying interactions between commodity and stock markets in Brazil. Overall, this dissertation contributes valuable insights into the nexus of commodities and economic uncertainty, shedding light on their complex interplay and implications for economic stability. By elucidating these relationships, this study offers practical guidance for policymakers, investors, and stakeholders, facilitating informed decision-making and enhancing risk management practices in an increasingly volatile global economy.

Keywords: Institutions. Agricultural Exports. Commodity Futures. Uncertainty

RESUMO

PINTO, Victor Henrique Lana, D.Sc., Universidade Federal de Viçosa, setembro de 2024. **Ensaios sobre commodities e incerteza econômica.** Orientador: Leonardo Bornacki de Mattos. Coorientadora: Lorena Vieira Costa Lelis.

Commodities desempenham papel fundamental no comércio global, incorporando matérias-primas e produtos agrícolas essenciais para crescimento e estabilidade econômica. Esta tese investiga a intricada relação entre commodities e incerteza macroeconômica, explorando como esses fatores interagem e influenciam a dinâmica econômica. Baseando-se em um arcabouço multidisciplinar, este estudo abrange dois artigos independentes que contribuem coletivamente para a compreensão dos mercados de commodities e incerteza econômica. O primeiro artigo examina o impacto das exportações de commodities na qualidade institucional em uma variedade de economias em desenvolvimento e avançadas de 1997 a 2022. Utilizando análise de dados em painel e a adesão da China à Organização Mundial do Comércio como um experimento natural, o estudo revela uma relação causal negativa significativa entre as exportações de commodities e o desenvolvimento institucional. Este achado destaca o efeito adverso da riqueza de commodities na eficácia institucional, destacando as complexidades da governança econômica em nações dependentes de recursos. No segundo artigo, o foco se volta para as interações entre futuros de commodities (especificamente boi gordo e soja) e o mercado de ações no Brasil de janeiro de 2014 a julho de 2023. Através de análise de séries temporais, investigamos a influência de mudanças de políticas nas correlações dinâmicas entre commodities e ações. Não observamos efeito economicamente significativo da incerteza econômica originada no Brasil e no exterior nas interações variáveis no tempo entre mercados de commodities e ações no Brasil. No geral, esta tese contribui com insights valiosos sobre o nexus de commodities e incerteza macroeconômica, lançando luz sobre sua complexa interação e implicações para a estabilidade econômica. Ao elucidar essas relações, este estudo oferece orientações práticas para formuladores de políticas, investidores e partes interessadas, facilitando a tomada de decisões informadas e aprimorando práticas de gestão de riscos em uma economia global cada vez mais volátil.

Palavras-chave: Instituições. Exportações Agrícolas. Futuros de Commodities. Incerteza

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1. INTRODUCTION

Commodities represent raw materials or primary agricultural products that can be bought, sold, and used as inputs in the production of goods and services. Commodities include agricultural varieties such as coffee and sugar; energy products like oil and gas; and metals such as gold, silver, and aluminum (Pimco, 2022). The importance of economic interactions involving commodities is multiple. On the one hand, the international trade of commodities allows countries to expand their markets, access different products and services, and promote economic growth (Taylor & Feenstra, 2020). On the other hand, standardized commodity contracts are negotiated on futures exchanges allowing producers and consumers to gain access to goods in a centralized and liquid marketplace (Hayes, 2021), and also, as a means for portfolio diversification (Tiwari et al., 2022). Whether commodities are distributed globally or traded in financial markets via futures contracts, they have become the central concern of various studies[1] for they can affect (or be affected by) economic uncertainty worldwide.

Economic uncertainty is perceived as the inability to predict possibilities associated with certain events (Franco, 2022; Keynes, 1936; Knight, 1921). Unexpected shocks create an unstable political and economic environment, which significantly affects international trade (Hassan et al., 2018) and financial markets (Boubakri et al., 2015). The rationale behind these associations is that uncertainty hinders bilateral negotiations as well as frightens off traders as prospective returns become relatively more volatile. In addition, measuring the true latent generating process of economic instability poses a challenge to empirical examinations and leads to the use of proxies (Baker et al., 2016; Jurado et al., 2015), such as the quality of institutions[2], and the uncertainty of the economic policy[3].

Institutional quality, for instance, includes how governments are selected, monitored, and replaced; the respect of citizens and the state for the institutions that

[1] For studies on commodity international trade see Bojnec and Ferto (2009), Mendonça et al. (2014), and Méon and Sekkat (2008). For papers on agricultural futures markets see Ouyang and Zhang (2020), Zhang and Yan (2020), and Zhu et al. (2021).
[2] The institutional quality indices are provided by the World Bank, namely, voice and accountability; political stability and absence of violence; government effectiveness; regulatory quality; rule of law; and control of corruption. These indicators merge the views of a great number of enterprises, citizens, and expert survey respondents in developed and developing economies (World Bank, 2021).
[3] The economic policy uncertainty index is provided by Baker et al. (2016). It uses text archives of major newspapers in a country and counts the number of articles containing policy-relevant terms. It refers to changes in policies that could affect the economic environment.

govern economic and social interactions; and the ability of the government to implement sound policies (World Bank, 2021). Institutions arise to reduce economic friction and soften uncertainty (Cabolis, 2020) and are determined by varying economic factors, namely, income per capita, international trade, and patterns of income distribution (Alonso et al., 2020). Alternatively, the uncertainty of the economic policy implies that governments are experiencing instability episodes, which decrease economic activity (Jackson et al., 2019), and affect the volatility of commodity futures returns (Watugala, 2019). Altogether, economic uncertainty either captured by institutional quality or the instability of the economic policy can not only influence, but also be affected by various economic factors, especially commodity exports (Liu & Chen, 2022) and the interactions between commodity futures and stocks traded in financial markets (Badshah et al., 2019)

The nexus between commodities and economic uncertainty is twofold. It can be seen, for example, as the connection running from the exports of commodities to institutions or also as the link between economic policy uncertainty and the interactions between the commodity and stock markets. The first of them considers that the specialization in commodity exports impacts the quality of institutions. That is, in resource-rich countries, who tend to export goods that use their abundant factor more intensively (Heckscher et al., 1991), ruling elites are often opposed to institutional enforcement as it imposes regular checks that hamper the appropriation of natural resource rents (Karl, 1997). As a result, the exports of commodities might represent a key factor in determining institutional quality.

The second of them takes into account that a high level of uncertainty concerning recurrent policy changes contributes to hightened volatility in financial markets. This means that unexpected shocks lead to a risk premium that is embedded in interest rates, making loans less affordable, which affects earnings projections, reducing stock prices and driving instability in financial markets (Badshah et al., 2019). For this reason, the interactions between the commodity and stock markets play a crucial role in asset allocation decisions and risk management (Badshah et al., 2019; Tiwari et al., 2022; Zhao & Wang, 2022), especially in a context of instability. In other words, the uncertainty of the economic policy could affect financial markets, and also, how they interact with commodity futures in several economies.

Economic shocks strike most, if not all countries. Increased uncertainty represents an early-warning signal of economic recessions affecting consumers, who need to cut back on spending; businesses, who resort to reducing production, investment, and employee compensation; and financial markets, which are usually more volatile with higher risk premium (Kupelian, 2017). In addition, commodities denote a vital element in both developing and advanced economies for they provide them with a larger set of agricultural goods via international trade, and also, because they constitute a means of protection against inflation and risk, and offer investors diversification through futures contracts.

Although commodities and economic uncertainty might have achieved prominence in every country in the world, it is probable that these variables have a contrasting effect on developing and developed nations. Taking, for instance, the association between commodity exports and institutional quality, it is important to bear in mind that less developed nations tend to be relatively natural resource-dependent economies (Khan et al., 2022), and that commodity wealth weakens the quality of institutions (Islam & Montenegro, 2002). Therefore, investigations regarding both developing and more advanced nations have the potential to deliver more accurate information on how primary trade could impact institutional quality.

Turning now to the connection running from economic policy changes to the interactions between commodity futures and stock prices, it is possible that such an association also affects several world economies. However, between richer and poorer nations, it is known that political instability is predominantly higher in emerging, developing democracies (Lupu & Riedl, 2013). In Brazil, for example, a sequence of corruption scandals such as the operation Car Wash[4], a serious fiscal crisis, and the presidential impeachment symbolize some events in the Brazilian government that could illustrate how uncertainty could influence the economic performance of the country (Ferreira et al., 2019). In other words, economic instability might alter investment decisions, and

[4] Car Wash or *Lava Jato* (in Portuguese) denotes one of the largest legal investigations on corruption in Brazil, bringing to light a set of illegal connections established between politicians, business corporations, and the public administration (Rios Petrarca, 2020).

determine asset allocation between the stock and commodity markets in a developing, more fragile economy, such as Brazil.

In this context, it is evident that there exist multiple links between commodities and economic uncertainty. For this reason, it is essential to try and understand the possible relationships between the exports of primary products and institutional quality as well as whether commodity-stock interactions are entangled with policy changes. Therefore, this doctoral dissertation consists of two papers on the relationships between commodities and economic uncertainty. The first of them aims to understand if the intensification of commodity exports weakens institutional quality in a set of developing and advanced economies from 1997 to 2022. The second paper questions whether recurrent policy changes (origininated in Brazil and abroad) affect the interactions between commodity futures (live cattle and soybeans) and the stock market in Brazil from January 2014 to July 2023.

Following this introduction, this dissertation presents the two papers that are independent in their focus and methodology; nevertheless, intertwined in their collective contribution to understanding the dynamics of commodities and economic uncertainty. Afterward, we present the conclusions of this doctoral dissertation, and lastly, a comprehensive list of references.

2. ARE COMMODITY EXPORTS A ROAD TO WEAKER INSTITUTIONS? CAUSAL INFERENCE THROUGH A NATURAL EXPERIMENT[5]

Victor Henrique Lana Pinto, Lorena Vieira Costa, Leonardo Bornacki de Mattos[6]

Abstract
Countries rich in natural resources are typically major exporters of commodities worldwide. Specialization in such goods may create incentives for certain groups to hinder a country's institutional development. Therefore, in this article we investigate the causal relationship between commodity exports and the institutional quality of 49 countries from 1997 to 2022. Given that commodity exports can both influence and be influenced by institutional quality, it is crucial to search for exogenous natural variations in commodity exports in combination with econometric modeling strategies. To address the endogeneity problem, we use a natural experiment – China's accession to the World Trade Organization – to establish a causal link between agricultural exports and institutional quality free from bias. Our results highlight that increased reliance on commodity exports is linked to detrimental effects on a country's institutions. It diminishes citizen participation, increases the risk of political violence, lowers public service quality, hampers private sector development, erodes trust in societal rules, and amplifies corruption. The findings confirm the negative relationship between commodity exports and institutional development. They also contribute to the resource-curse debate by showing that agricultural exporters with abundant natural resources negatively impact institutions. Additionally, a causal relationship is established, demonstrating that focusing on commodity exports reduces institutional effectiveness, with the study's quality enhanced by using various commodity product bundles and alternating institutional quality measures.

Keywords: Commodity Exports; Institutions; Natural Experiment; Two-Stage Least Squares

JEL codes: F13; F55; Q0

2.1. Introduction

Various researchers have long investigated the determinants of the institutional quality of countries worldwide since it plays a central role in defining the long-term growth prospect of a nation. Economies with higher levels of institutional quality (i.e. voice and accountability; political stability and absence of violence; government effectiveness;

[5] This study was financed in part by the Coordenação de Aperfeiçoamento de Pessoal de Nível Superior - Brasil (CAPES) – Finance Code 001 – and by the Conselho Nacional de Desenvolvimento Científico e Tecnológico (CNPq). This paper was published in *Agribusiness: An International Journal* on Apr 22, 2024 and is available at https://doi.org/10.1002/agr.21943
[6] The author gratefully acknowledges the financial support received from the CNPq through a PQ-2 research productivity grant.

regulatory quality; rule of law; and control of corruption)[7] attract more foreign investment and use human and physical capital more efficiently, resulting in better growth indicators in the long run (Acemoglu et al., 2005; Qiang & Jian, 2020; Sala-i-Martin & Subramanian, 2013).

In light of these findings, some studies examine what determines different aspects of institutions. Acemoglu et al. (2001), for instance, consider the effects of climate and colonization policies to explain the quality of institutions and subsequent growth of 64 former European colonies in various regions of the world. These authors explain that institutions today are not predetermined by colonial policies; however, these researchers emphasize that the colonial experience is indeed one of the many factors affecting institutions. Similarly, Chong & Zanforlin (2000) analyze the impact of colonial heritage on institutional development for 101 nations worldwide and unveil that countries with a common law tradition show a positive association with institutional quality; that is, the legal system inherited from colonization passes onto the institutions of a country.

Wei (2000) offers an interpretation of the connection between trade openness and good governance. The author conducts an analysis for more than 160 countries and finds that naturally more open economies do exhibit lower corruption levels. In other words, foreign trade could, on average, affect a country's corruption level, i.e., its institutions. More recently, Javaid et al. (2017) investigate the association between income, tax collection, military spending, openness, and indebtedness and the quality of institutions for a set of 11 developing Asian countries. These authors uncover that, except for military expenditure and openness, those variables significantly determine the level of institutional quality for those economies.

According to Lehne et al. (2014), rule of law and control of corruption have shown to be better in countries that are more open to trade, financial flows, investments, and that do not have significant natural resource endowments. These authors, similarly to Islam and Montenegro (2002), have also observed a negative correlation between commodity wealth and the quality of institutions for more than 130 economies.

Although the prominence of countries' institutional quality has been vastly

[7] The aforementioned indicators refer to the Worldwide Governance Indicators provided by the World Bank. These aggregate indicators merge the views of a great number of enterprises, citizens, and expert survey respondents in developed and developing economies (World Bank, 2021).

recognized, its determinants are still quite challenging to disclose. A relevant question from a trade policy perspective is whether the specialization in commodity exports affects the quality of institutions. Larsen (2006) observes that oil discovery and exports, for instance, have positively affected the economies of more advanced nations. For this reason, one could conclude that such a pattern would be similarly advantageous for developing countries. Nevertheless, various studies have argued whether the abundance of natural resources in those economies would be a blessing rather than a curse (Ablo, 2015; Atkinson & Hamilton, 2003; Kopinski et al., 2013; Papyrakis & Gerlagh, 2004; Satti et al., 2014). Literature on the natural resource-curse hypothesis shows that less developed countries that are rich in natural resources are more inclined to revenue mismanagement of commodity goods, poor governance (Watts, 2004), corruption (Roberts, 2015), and socioeconomic and political crisis (Billon, 2006) that can weaken democracy (McFerson, 2010), stability, growth, and development (Mehlum et al., 2006; Robinson et al., 2006). In other words, as stated in (Lehne et al., 2014) the abundance of natural resources relative to the size of the economy negatively affects the institutional development of a given nation.

Guriev et al. (2010) note that the performance of commodity-rich countries tends to gradually make institutions less effective. Theoretically, the rationale behind the channel linking commodity exports and institutions is that in resource-rich nations ruling elites are often opposed to institutional enforcement as it imposes regular checks that hinder the appropriation of natural resource rents (Karl, 1997). Hodler (2006) explains that natural resources cause fighting activities between rivalling groups, which reduces production and weakens property rights, making productive activities less attractive. In case the number of rivalling groups is sufficiently large, the author emphasizes that the aggregate production decrease exceeds the natural resources rent effect. Additionally, in line with the trade literature, the question of whether commodity exports affect the quality of institutions is based on the Heckscher-Ohlin model that implies that a country will export goods that use its abundant factor intensively.

In this paper, we investigate the causal effect of commodity exports on the

institutional quality of a set of 49 countries[8] between 1997 and 2022, and seek to answer whether variations in agricultural exports impact a country's institutional level. Our sample includes 26 developed and 23 developing exporters of commodities (aggregation of agricultural products, livestock and meat, energy products, and metals). In the past decades, various emerging nations have developed their exporting basket into becoming exporters of commodity goods (World Integrated Trade Solution, 2023). Therefore, our panel data include not only advanced countries, but also a number of developing and transition economies, which is of particular relevance to this study.

In this context, we expect that countries that are rich in natural resources are those exporting mainly commodity goods worldwide (Heckscher et al., 1991), and that specialization in commodities might result in an incentive for certain groups to hinder a country's institutional development (Guriev et al., 2010; Karl, 1997; Lehne et al., 2014). For this reason, our hypothesis is that commodity trade causes weaker institutions, mainly in developing countries.

Since commodity exports might, at the same time, cause and be affected by institutional quality, it is crucial to search for exogenous natural variations in commodity exports in combination with econometric modeling strategies. This paper exploits a significant change in global trade dynamics to evaluate the relationship running from commodity exports to the quality of a country's institutions. In 2001, China finally gained accession to the World Trade Organization (WTO) after a long period of negotiations. China's accession to the WTO has impacted both the Chinese and the world economy since 2001. It resulted in staggering growth of exports, and a reduction in import tariffs into China. Boden (2012) highlights that the accession demands placed upon China brought the nation into a period of intensive trade liberalization and weakened state-run businesses. Furthermore, as of 2001, more economies have benefited from expanded access to the Chinese market, and vice-versa.

The identification strategy in this paper uses the fact that the accession of China to the WTO provided exogenous variations in worldwide trade. The dramatic reduction in

[8] Developed: Australia, Austria, Bulgaria, Canada, Croatia, Finland, France, Germany, Greece, Hungary, Ireland, Italy, Japan, Latvia, Lithuania, the Netherlands, New Zealand, Norway, Poland, Portugal, Romania, Slovak Republic, Slovenia, Switzerland, United Kingdom, and the United States. Developing: Argentina, Brazil, Chile, Colombia, Costa Rica, Guatemala, India, Indonesia, Israel, Korea, Madagascar, Malaysia, Mauritius, Mexico, Nicaragua, North Macedonia, Russian Federation, Senegal, Singapore, South Africa, Turkey, Uruguay, and Zambia.

import tariffs as well as in tariff rate quota[9] have played a key role in promoting freer trade into China. The implications of China's accession to the WTO have received close attention and caused widespread speculation mainly in respect of its domestic agricultural sector (Bingshen & Wan, 2003). The rationale behind these concerns relate to the prospect increase of foreign commodity goods into the Chinese market as a result of the concessions made by the Chinese government to gain WTO membership status. Thus, we use this exogenous source of variation in trade to estimate the impact of commodity exports on institutional quality.

The contributions of this article are manifold. First, it assesses whether the 'natural resource – weak institutions' nexus also holds for the relationship between commodity exports and institutional quality. This represents to the related literature a novel approach towards the understanding of the natural resource-curse hypothesis. Second, it focuses on a more disaggregated analysis. That is, it evaluates how different types of commodities exports (animal, vegetable, food products, minerals, and fuels) impact institutional quality, which improves the quality of its empirical analyses and denotes a useful supplement to existing studies. Third, this article accounts for the income structure of the sample and how it can affect institutions.

Another contribution is our identification strategy. We use China's accession to the WTO as a natural experiment to deal with endogeneity, and explore the impact of commodity exports on institutional quality free from any bias. The use of this approach provides some distinct advantages. Firstly, by harnessing exogenous variation inherent in real-world events, such as China's accession to the WTO, we can effectively circumvent concerns regarding endogeneity. This not only bolsters the credibility of our findings but also enhances confidence in the causal relationships we uncover. Additionally, as natural experiments occur organically, our approach alleviates concerns about the manipulation of variables, which ensures the integrity of this research. Overall, our reliance on a natural experiment as an identification strategy not only strengthens the robustness of our findings but also underscores the relevance and practicality of our research in informing policy and furthering understanding in our field.

[9] Tariff rate quotas (TRQs) permit a pre-determined quantity of a product to be imported at lower import duty rates (in-quota duty) than the duty rate normally available for that good (European Comission, 2021).

Finally, policymakers aiming at institutional development might benefit from the understanding of the possible impact between commodity trade and institutions, especially in resource-endowed economies. In addition, using various institutional dimensions offers knowledge to legislators as to how commodity exports influence institutions, i.e. via an impact on a country's regulatory quality, rule of law, or control of corruption, for instance. Still, the results of this research can be used in favor of government efforts to diversify a country's exporting agenda, and consequently, produce an effect onto institutional quality. That is, if the international trade of commodities affects institutions, countries could focus on the production (and trade) of other varieties. This, in turn, could lead to the substitution of a commodity based exporting agenda for a less resource-dependent type of trade.

The remainder of this paper is organized as follows. Section 2.2 elaborates on the conceptual mode. In section 2.3, we present the identification strategy and discuss our exogenous source of variation in commodity trade. In section 2.4, we describe the data on institutional quality and commodity exports. Section 2.5 introduces our empirical strategy. In section 2.6, we present the causal effect of commodity exports on institutions. Section 2.7 contains our conclusions and in section 2.8, the appendix.

2.2. Conceptual model

This section relies on the model of Mehlum et al. (2006) and lays out the theoretical basis linking commodity exports (via the natural resources abundance channel) and institutional quality.

The model in Mehlum et al. (2006) considers the endowment of natural resources and not the exports of primary goods. Based on the assumptions of the Heckscher-Ohlin model of trade (H-O model)[10], countries with more natural resources are those that most trade commodities. This model implies that countries that are rich in natural resources (abundant in land) tend to produce and export products that use this factor of production more intensively, as is the case of commodities. Thus, this study uses not only the H-O model, but also the theoretical model of Mehlum et al. (2006) to determine the association between the exports of commodities and the quality of institutions.

[10] See Heckscher et al. (1991) for more details of the H-O model.

Mehlum et al. (2006) provide a conceptual model aiming to relate the extent to which resource rents appropriation depends on the quality of institutions. Alternatively, paper aims to unveil the opposing association; that is, the connection running from primary exports (as a result of natural resource endowment) to institutional quality. However, at the same time that institutions could facilitate or hinder resource rents (Mehlum et al., 2006), resource abundance might also lead to a deterioration of institutional quality (Rosa & looty, 2012). For this reason, this subsection casts light on the theoretical relationship between those variables regardless of the direction of this effect, which will be dealt with in section 2.3.

Formally, Mehlum et al. (2006) make a distinction between producer- and grabber-friendly institutions. In the former type, production and rent-seeking activities are complementary yet in the latter kind production and rent-seeking actions compete with one another. In other words, the theoretical model in Mehlum et al. (2006) focuses on the allocation of entrepreneurs between production and unproductive rent extraction (grabbing). The authors indicate that with grabber-friendly institutions there exist gains from the specialization in unproductive activities, such as resource exploration, due to a weak rule of law, corruption, or malfunctioning bureaucracy, for instance. In their model, the total number of entrepreneurs is $N = n_P + n_G$, where n_P and n_G are producers and grabbers, correspondingly. In this study, the main focus is the fact that grabbers target rents from natural resources R and use all their capacity to take as much as possible of this rent. Theoretically, the parameter λ captures institutional quality. When it equals zero, the institutional arrangement is weak and the system is completely grabber-friendly, i.e.: ruling elites are able to extract the entire rent, each of them obtaining R/n_G. Naturally, when $\lambda = 1$, there are no gains from specialization in primary goods as both grabbers and producers each get the share R/N of resources. The payoff π_G to each grabber is the factor *s* times R/N, as follows:

$$\pi_G = s\,R/N \tag{1}$$

while the share of the resource rent of each producer is $\lambda s\,R/N$. When countries have good institutions ($\lambda \to 1$), the factor *s* is decreasing.

The sum of the shares of the resource rents for each group, grabbers and producers, cannot exceed one. For this reason, the following constraint must hold[11]:

$$(1-\alpha)s + \alpha\lambda s \leq 1 \quad (2)$$

where $\alpha = n_P/N$ means the fraction of producers. Solving for s in the condition (2), it follows

$$s = s(\alpha, \lambda) \equiv \frac{1}{(1-\alpha)+\alpha\lambda} \quad (3)$$

where the factor s is an increasing function of α and a decreasing function of the institutional quality λ. This means that the factor *s* increases as the number of producers gets higher, and as institutions become weaker.

The profits of a producer π_P represents the sum of profits from production π and the share of resource rents:

$$\pi_P = \pi + \lambda s(\alpha, \lambda)\, R/N \quad (4)$$

The authors elaborate more on their model to also determine π. Since they are specially interested in how resources affect growth, and for that, associate resources to a country's institutional framework, following Murphy et al. (1989) they embed their mechanism in a development model with joint economies in modernization[12]. After a set of assumptions and calculations, Mehlum et al. (2006) verify that $\pi = \pi(n_P)$ and that $n_P = \alpha N$, i.e.: $\pi = \pi(\alpha N)$:

$$\pi_P = \pi(\alpha N) + \lambda\pi_G \quad (5)$$

Condition (5) looks at the equilibrium allocation of entrepreneurs, between producers and grabbers. This means the relative profits of the two activities from (1) and (4), where both functions π_G and π_P are increasing in the fraction of producers α. Finally, the authors show that the economy might then be in one of the following types of equilibria:

a. Production equilibrium: in this case scenario profits of producers are greater than or equal to profits of grabbers and all entrepreneurs are producers; i.e.: $\pi_P \geq \pi_G$ and $\alpha = 1$.

[11] Mehlum et al. (2006) note that when rents are not wasted in the sharing, condition (2) must hold with equality.
[12] For details of the algebra used to disclose the profits from production see Mehlum et al. (2006). For the entire formalization of the model see Murphy et al. (1989).

b. Grabber equilibrium: in this case scenario profits of producers are equal to profits of grabbers while some entrepreneurs are producers and some others are grabbers; i.e.: $\pi_P = \pi_G$ and $\alpha \in (0,1)$.

In sum, this subsection attempts to demonstrate the interactions between resource abundance, export performance and its effects onto institutional quality. The H-O model of trade serves to establish a relationship between resource abundance and primary exports. In addition, the model of Mehlum et al. (2006) shows through conditions (1) and (4) that the payoffs of both producers and grabbers respond to variations in the factor *s* which, in turn, is a function of the quality of institutions λ. In other words, natural resource endowments might determine the export concentration of primary goods and these products are somehow intertwined with institutions, regardless of the direction of this effect. Section 2.4 discusses this further.

2.3. Identification strategy

One natural question that arises is the cause behind fluctuations in commodity exports among countries over time. Are these variations a result of actions taken by the governments of these sampled countries? If so, are these actions linked to unobserved characteristics of these governments that might also impact institutional quality? Alternatively, could these variations in commodity trade be attributed to an increase in foreign demand? If the latter scenario holds true, there is less cause for concern, as it suggests an external source of variation. However, when considering the first two possibilities, it is essential to bear in mind that these fluctuations in exports among countries over time may be driven by the government's strategic decisions, actions aimed at achieving a positive trade balance, the use of domestic agricultural subsidies, and specific investments to enhance production capacity, among other factors. In other words, if we observe any of these two possibilities, we must consider that our variable of interest could potentially be endogenous. This is because these unobserved aspects (government actions) are also connected to the quality of institutions. If we fail to account for them, our estimates would not only capture the causal effect of commodity trade on institutional quality but also the bias stemming from these unobservable factors.

Funke and Holly (1992) argue that several previous studies have placed emphasis

on demand factors. For these researchers, such models have generally been rather unsuccessful in explaining long-run trends in export performance. Majeed and Ahmad (2006) investigate the determinants of exports in developing economies and observe that the supply side factors are much more important for explaining export performance than demand side factors. These authors observe that private domestic investments, for example, represent a permanent and reliable channel to enhance production capacity, and consequently, improve export performance. These perceptions give us an early indication of the potential endogeneity of our variable of interest.

Similary to Majeed and Ahmad (2006), Briones and Rakotoarisoa (2013) notice that in developing countries variations in agricultural trade flows are subject to supply and demand drivers. For them, supply side factors involve domestic production and processing technology and logistical systems. From a foreign demand perspective, the authors point out that increases in urban populations, with consumer preferences oriented towards food safety, high quality, and differentiated goods, tend to favor production and distribution systems.

Again, in our study, domestic supply side factors are our main concern for they might be endogenous, i.e.: they may be the result of governmental own actions aiming to encourage foreign trade. These actions can be correlated with unobservables that could also affect institutional quality. Hence, we take into account the main findings in the aforesaid studies and rely on an exogenous source of variation in our country-level commodity exports. For that, we use China's accession to the WTO in 2001 as a natural experiment to study the relationship between commodity trade and the quality of a country's institutions.

On November 11, 2001, China acceded to the WTO becoming its 143rd member. The consequences of China's accession to the WTO have been given substantial attention in the past two decades because of the concessions made by Chinese negotiators and their implications for trade and domestic policy changes. One particular issue has emerged in the context of the Chinese agricultural sector. The Chinese exporting agenda was mostly concentrated on labor-intensive manufactured goods due to the abundance of inexpensive labor (Chen, 2009). This focus on labor-intensive manufacturing shifted workers away from the commodity to the secondary sector (Martí

et al., 2011) and could be interpreted as a better allocation of production factors caused by China's accession to the WTO.

Chen (2009) stresses that the accession of China to the WTO means that the country's imports would all face the same tariffs and barriers to trade, thus making each country's commerce with China a reflex of its comparative advantages, not its political power. Bingshen and Wan (2003) notice that China agreed to cut import tariffs for fruit, meat, dairy products, and other agricultural commodities from 21 to 15 percent by 2004. For varieties such as grains, cotton, edible oil, sugar, and wool, the tariff rate quota regime was applied and believed to have a greater impact on China's agricultural sector than its tariff reduction commitment (Bingshen & Wan, 2003).

Hence, what we observe is that perhaps the concessions China had to make during its accession to the WTO may have facilitated easier access for foreign commodity products to enter the Chinese market. The reallocation of Chinese workers from the commodity to the secondary sector, as a potential consequence of the accession of China to the WTO, is significant because it may have resulted in an increased Chinese demand for commodity goods to compensate for this workforce shift. If this hypothesis holds true, the potential rise in Chinese demand for commodity goods, as a response to the workforce shift, might have externally influenced the commodity exports of numerous commodity-exporting countries globally, particularly in developing and transition economies.

It is true that countries engaged in some early negotiations with China prior to its accession to the WTO, and some may have anticipated potential benefits from closer economic ties with China. Nonetheless, the full extent of the benefits and the rapid transformation that would follow China's accession might not have been fully anticipated. While negotiations were ongoing, countries likely expected increased access to China's market and opportunities for their own exporters. However, the scale and speed of China's integration into the global economy, as well as the profound changes it would bring to trade dynamics, were not fully foreseen. For example, the extent to which China's production would shift from the primary to the secondary sector and the resulting impact on global trade patterns could not have been perfectly predicted. Therefore, although countries may have had some expectations regarding the benefits of China's accession, the exogenous shock aspect still holds because the full magnitude and implications of

China's integration into the WTO were not entirely foreseeable. In that sense, the rapid and transformative nature of China's accession and subsequent impact on global trade patterns can be regarded as an exogenous shock.

In summary, our identification strategy relies on using China's accession to the WTO as a natural experiment to generate exogenous variation in the commodity exports of our selected countries and establish a causal link between these exports and their respective institutions. As China stands as one of the foremost global trading nations (World Integrated Trade Solution, 2023), its agricultural imports have the potential to heavily influence the international demand for various commodities, particularly in developing and transitioning economies who are often reliant on natural resources. Therefore, we argue that Chinese imports of primary-based products effectively reflect the external factors that shape a country's decisions regarding its policies on the exports of commodities.

To shed some light on the association between the accession of China to the WTO in 2001 and commodity exports (first-stage relationship), Figure 1 shows the aggregate trade of commodity goods exported by our sampled countries to the world from 1997 to 2022.

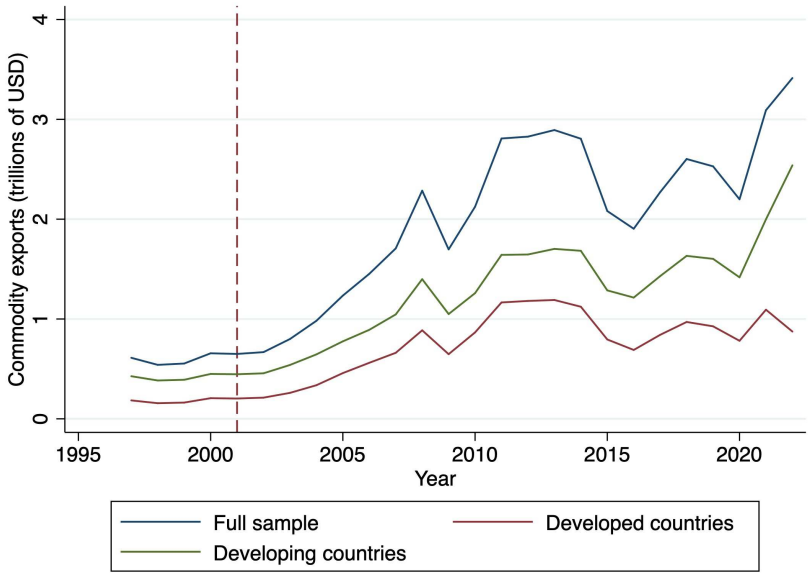

Figure 1. Aggregate commodity exports for full sample, developed, and developing countries between 1997 and 2022.
Source: Own calculations.

Figure 1 demonstrates that, as of 2001, the aggregate exports of commodity varieties followed a consistently increasing pattern until the year 2008 for the entire sample, as well as for the subset of developed and developing economies. This surge indicates a dramatic increase in the shipment of commodity goods from the sampled countries worldwide. Figure 1 suggests that developed countries seem to be trading higher volumes of commodity goods in comparison to less advanced economies. Throughout the years, these trends, however, seem to follow a somewhat similar pattern.

From 2008 to 2009, there was a reduction in exports shown in figure 1, possibly related to the subprime crisis that hit the global economy during those years. A rapid recovery period stands out between 2010 and 2011, followed by a positive, yet modest performance until 2013. Figure 1 also exhibits a sharp decline between 2013 and 2016, a period marked by significant global events that could have influenced the performance of commodity trade, including the Chinese crash, the Organization of Petroleum Exporting Countries (OPEC) cut in crude oil production, the trade policies of former US President Donald Trump, and Brexit. Lastly, figure 1 shows the possible effects of the Covid-19 pandemic on the international trade of commodity varieties, particularly in 2020, when it had a significant impact.

While it is not possible to draw strong conclusions from figure 1 regarding the relationship between China's accession to the WTO in 2001 and the increase in commodity trade, it is evident that, starting from that year (at least until 2008)[13], a greater volume of commodity varieties was exported by the sampled countries. Additionally, it should be noted that the accession of China to the WTO resulted in a shift of workers into the manufacturing sector, increasing China's demand for commodity goods sourced from abroad. Consequently, based on an examination of the raw data in figure 1, a positive correlation may exist between China's WTO membership and the exports of commodity varieties. In summary, in our identification strategy we use this observation as a potential source of exogenous variations in commodity trade to determine its causal effect on institutions.

[13] The subprime crisis in early 2008 caused a sharp decline in the world's economic activity and liquidity, which mirrors the pattern observed in figure 1 up to late 2009.

2.4. Data

The data used in this paper come from multiple sources. Our panel gathers information at the country-level between 1997 and 2022[14] for 49 nations. Institutional quality comes from the World Governance Indicator (WGI), organized into six dimensions, namely, voice and accountability; political stability and absence of violence; government effectiveness; regulatory quality; rule of law; and control of corruption. These data are reported in standard normal units of the governance indicators ranging from -2.5 to 2.5 (higher figures indicate higher levels of governance). In line with Chong and Calderón (2000) and Tebaldi and Elmslie (2013)[15], we also compute a combined measure of institutional quality based on the annual simple average of all six variables for each of our sampled economies.

Figures on commodity trade flows are from the World Integrated Trade Solutions (WITS). We use the 2-digit Harmonized System (HS) classification of merchandise to segregate commodity from non-commodity varieties. More explicitly, we rely on the range of products between the codes 01 and 27[16], comprising the sum of several types of commodity goods shipped from each of our set of countries to the world.

Finally, our control variables come from the World Bank Development Indicators and CEPII. The sample is comprised of 49 economies. These countries have been randomly selected through the following data processing: First, we collected information on commodity trade from all reporting countries between 1997 and 2022; next, we summed up countries' information at the product level and have kept solely those nations who were exporters of commodity goods throughout the entire 1997-2022 timespan. Then, we organized the data on institutions and merged with our trade observations. Lastly, we obtained our controls for the matching countries and dropped the observations that were missing important variables and/or only existed for one period.

[14] We note missing values for 1997, 1999, 2001, and 2022 for all countries in our dataset. We followed Topal and Sahin (2017), who employed linear interpolation, and covered this issue accordingly. In our study, this method creates a linear interpolation of institutions on commodity exports for the missing values of the six dimensions of institutional quality.

[15] Chong and Calderón (2000) use data from the International Country Risk Guide (ICRG) and the Business Environmental Risk Intelligence (BERI) to compute an institutional quality index, which is the simple average of five institutional dimensions. Tebaldi and Elmslie (2013) build a measure of institutional quality based on the simple average of all variables produced by Kaufmann et al. (2010) in the WGI data.

[16] This range of goods includes animal, vegetable, food products, minerals, and fuels.

2.5. Empirical strategy

2.5.1. Econometric model

We set the econometric model as follows:

$$Inst_{it} = \alpha_0 + \alpha_1 lnX_{it} + \sum \beta V + \varepsilon_{it} \quad (5)$$

$Inst_{it}$, the dependent variable, represents the institutional quality of country *i* in year *t*. We run six separate models[17] using each of the six dimensions of institutions presented in section 2.4, and an alternative estimation in which our outcome variable represents the average institutional quality index of country *i* in year *t*. Our variable of interest is lnX_{it}, which denotes the natural logarithm of the aggregate commodity exports (in US dollars) of country *i* in *t*. The key coefficient α_1 indicates the effect of variations in commodity exports on institutions. The specialization in commodity goods might affect the quality of institutions negatively for it can create a less dynamic environment, which fuels a lower demand for sound institutions (Guriev et al., 2010). Furthermore, ruling elites might (especially in developing countries) hinder competition, promote corruption and nepotism (Alonso & Garcimartín, 2009; Hodler, 2006; Karl, 1997).

All our remaining explanatory variables were also selected on the basis of their theoretical relation to the quality of institutions. We mostly follow Alonso and Garcimartín (2009) and add vector V, which represents a series of predetermined controls, including the total natural resources rents and the natural logarithm of *per capita* GDP of country *i* in year *t*. We also control for the geographic location of country *i* and whether it is landlocked or not. We provide a brief theoretical rationale to these choices. First, we ought to consider that natural resources can lower institutional quality by fostering rent-seeking activities and replacing tax revenues by other less transparent revenue sources Alonso and Garcimartín (2009). Second, theory suggests the existence of a positive association between *per capita* GDP and the quality of institutions since higher income determines the availability of resources to build good institutions (Islam & Montenegro, 2002). Lastly, institutions can also be impacted by geographical conditions. A country's location near the tropics, and its lack of access to the sea could, for example, influence the development of strong institutions (Easterly & Levine, 2003; Gallup et al., 1998).

[17] In our study, we conducted each estimation separately rather than simultaneously. This approach alleviated concerns about statistical size distortions, thereby eliminating the need for Bonferroni correction in our regressions.

2.5.2. Framework to identify causal effect

Based on equation (5), we are far from capturing the causal effect of commodity exports on institutional quality for we are probably involved in endogeneity. The governmental strategies adopted by each nation to develop or boost its trade policy and increase its institutional quality might be simultaneously determined by unobservable aspects, such as the motivation of the country leader, omitted in the error term. It is hard to identify the causal effect of commodity exports on institutions because of unmeasurable country factors, such as its leader's motivation that may change in time as leaders alternate office, for instance.

In this paper, we manage the endogeneity problem by using the instrumental variable (IV) approach through the two stage least squares estimator (2SLS) as in Autor et al. (2013), and Benguria and Ederington (2023). The conduct of particular agricultural trade policies and the use of sector-specific subsidies are some examples of how the government of a country may choose and manipulate its volume of commodity exports. However, if the foreign demand increases substantially, domestic producers might as well feel encouraged to ship their output overseas, and then, the changes observed in the commodity trade figures should be exogenous.

The construction of the IV follows[18]: the natural logarithm of the aggregate Chinese imports of commodity goods, denoted as $lnchinaM_t$. Since China represents one of the largest global traders in the world (World Integrated Trade Solution, 2023), the variable $lnchinaM_t$ vastly determines the foreign demand for commodity varieties to many countries, especially developing and transition nations who tend to be natural resource-dependent economies. Hence, we believe our instrument captures the exogenous variation that affects the country's decision on the conduct of commodity trade policies. In other words, $lnchinaM_t$ affects the endogenous variable lnX_{it} (first-stage regression). We also believe that our instrument has no power over our outcome variables other than through the first-stage channel. It is unlikely to imagine that the Chinese imports of commodity goods could produce a direct effect on the institutional development of countries. In other words, the volume of commodity goods imported by China is not an

[18] Similar instrumental variable approaches have been used in Autor et al. (2013), Benguria and Ederington (2023), and Costa et al. (2016).

indication of any factors that can cause a difference in institutional quality, other than the effect on commodity exports. Hence, we consider that our instrument is as good as randomly assigned or independent of potential outcomes, conditional on our predetermined set of covariates V.

Furthermore, our instrument seems uncorrelated with (at least some) unmeasurable country-specific factors. As discussed in Chen (2009), the accession of China to the WTO indicates that its imports would all be subject to the very same tariffs to trade. Then, each country's trade with China would solely regard its comparative advantages, not its political (unobserved) influence. Under the WTO agreements, countries cannot discriminate among their trading partners or grant special favors, for example (WTO, 2021). Hence, we argue for validity of our instrument that China's accession to the WTO is uncorrelated with exporting country-specific unobservables in ε_{it}. More clearly, let us say that a country leader is strongly motivated to develop their exports. Before China's membership in the WTO, this leader could, for instance, try and establish a closer relationship with China different from their conduct towards other trade partners. However, under the WTO agreement, this unobserved factor can no longer affect $lnchinaM_t$, i.e. it is exogenous[19].

The first- and second-stage econometric models follow:

$$lnX_{it} = \lambda_0 + \lambda_1 lnchinaM_t + \sum \eta V + \upsilon_{it} \quad (6)$$

$$Inst_{it} = \alpha_0 + \alpha_1 ln\hat{X}_{it} + \sum \beta V + \varepsilon_{it} \quad (7)$$

In sum, we have a causal effect if two conditions are satisfied: Chinese imports are a strong determinant of commodity exports and this is the only channel through which they are correlated with institutional quality of our countries.

[19] A reviewer suggested considering an alternative instrumental variable in our estimations – the initiation of negotiations between China and the WTO. In our effort to incorporate the reviewer's comment, we noted that negotiations aiming at China's accession to the WTO began in 1986 (WTO, 2001), which would encompass a period beyond the one we consider in our study. Additionally, we recognized that this variable would be the result of China's own decision to join the WTO, potentially violating the exclusion restriction. The reviewer also pointed out that it is possible for anticipatory effects to occur, rendering our instrument not fully exogenous since China's accession could be correlated with the timing negotiations started. In such cases, a "time" effect could be nested. We followed their suggestion and also considered the year China acceded to the WTO as a dummy to account for potential temporal effects. We estimated all our equations, including a dummy variable for the year China acceded to the WTO. However, we observed that the dummy variable is highly correlated (corr.: 0.71) with our instrument. Therefore, incorporating such a dummy variable in our estimation would violate the independence assumption, as our instrument should not be correlated with other regressors included in the main equations.

2.6. Results

2.6.1. Descriptive statistics

In order to compare whether the group of countries that export relatively more commodities are similar or different from those that export fewer products of this type, we use certain criteria to define the export intensity of these goods. Specifically, we adopt two approaches. First, we identify those countries that are net exporters of commodities. This refers to countries whose exports of agricultural goods exceed their exports of non-agricultural products. Positive outcomes indicate that the country exports a greater proportion of commodities compared to manufactured products, making it a net exporter of commodity goods. Second, we compute the percentage of agricultural exports in relation to each country's total exports. After that, we consider those countries whose share of agricultural exports in respect of their total exports is equal to or greater than 60%.

After establishing these criteria, we obtain the mean values of all variables included in our empirical specification and other covariates (namely the Gini index, arable land, CO_2 emissions, fertilizer consumption, and R&D expenditure[20]) for each group aiming to characterize our sample and discuss their differences (if any). In addition, we perform a mean-comparison test to verify if these two groups are statistically similar in those characteristics. Table 1 shows these results.

Table 1 – Mean values of variables for countries that are more (or less) intensive in the exports of commodity goods

Variables	Net exporter of commodities?		Higher share of agricultural exports?	
	Yes	No	Yes	No
Voice and Accountability	0.58***	0.75	0.59***	0.74
Political Stability and Absence of Violence	0.09*	0.66	0.10^{NS}	0.60
Government Effectiveness	0.62*	0.83	0.71^{NS}	0.79
Regulatory Quality	0.64***	0.84	0.67**	0.82
Rule of Law	0.43***	0.77	0.49***	0.72

[20] These covariates were solely included in the descriptive statistics analysis. They are meant to help characterize the sample and since they could potentially yield econometric issues related to endogeneity they were not included in equations (6) and (7).

Control of Corruption	0.40***	0.69	0.46**	0.65
Average institutional index	0.46***	0.76	0.51**	0.72
Natural resources rent (% of GDP)	4.09***	1.77	4.58***	1.81
Latitude (degrees)	12^{NS}	14	15^{NS}	13
Landlocked (1 = yes; 0 = no)	0.07***	0.15	0.09**	0.13
GDP per capita (USD)	25907.34^{NS}	25393.77	26891.51^{NS}	25111.41
Gini (0% = perfect equality)	39.15***	36.24	39.6***	36.5
Arable land (ha per person)	0.35***	0.25	0.35***	0.26
CO_2 emissions (metric tons per capita)	6.9***	5.7	7.3***	5.7
Fertilizer consumption (kg/ha of arable land)	273.48*	413.65	293.67^{NS}	395.67
R&D expenditure (% of GDP)	1.05***	1.43	1.11***	1.38
Observations	390	876	309	957

Notes: ***, **, and * means are statistically different from the 'control' group at 1%, 5%, and 10%, respectively. NS means are statistically equal to the 'control' group. Institutional measures range from -2.5 to 2.5 (higher figures indicate higher levels of governance). Figures rounded to two decimal places.
Source: Own calculations.

Table 1 considers six institutional indicators as well as the average institutional index, and the remaining controls, namely natural resources rents, location, access to the sea, and GDP *per capita*. In addition, table 1 also presents mean values for other characteristics that could provide some interesting insights into both groups of countries. By doing this, table 1 allows to verify whether countries that are more (or less) intensive in the exports of agricultural varieties have similar levels of inequality, arable land, CO_2 emission, fertilizer consumption, and R&D expenditure.

Columns 1 and 3 show the mean values of these variables for those countries whose exports are more concentrated on commodity products (net exporters of commodities and those with a higher share of agricultural exports). Columns 2 and 4 present the average of all variables for those economies whose exports are less focused on agricultural goods. The results we obtain using both criteria (net exporter and share of agricultural exports) are quite similar. Therefore, the characterization of our sample does not vary widely when we consider a different approach to identifying those economies whose exports are intensive in commodities.

We observe in table 1 that countries with a higher intensity of commodity exports tend to have lower institutional figures. When we look at those countries who are net exporters of commodities, we note that the mean value of Government Effectiveness, for instance, is 0.62 while the average institutional index equals 0.46. table 1 shows that Political Stability seems to be, on average, fragile in countries concentrated on the exports of commodity goods. Moreover, we observe that the index for Rule of Law denotes that individuals and firms have little confidence and not always abide by the rules of the society in those economies. A similar pattern follows for the remaining institutional measures.

The results in column 1 of table 1 reveal that the group of countries with a higher intensity of agricultural exports present, on average, 4.09% of their GDP in natural resources rents. Furthermore, the mean values of latitude for this group suggests that their relative location near the equator is not statistically different between groups. The information reported in column 1 also indicates that the countries in this group are not landlocked, which could facilitate foreign trade interactions. Lastly, we note that the average value of GDP *per capita* is not statistically different from zero between those countries whose exports are more and less focused on commodities.

When we look at the group of countries with a lower intensity of commodity exports, we notice that they present higher figures for all institutional indicators. For example, the economies in this group show, on average, an index of 0.84 (column 2) and 0.82 (column 4) for Regulatory Quality, denoting the ability of the government to formulate and implement sound policies and regulations. Besides, countries more focused on the exports of non-agricultural goods present an index of 0.69 (column 2) and 0.65 (column 4) for the Control of Corruption indicator meaning that in these societies public power cannot be vastly exercised for private gain. We do not observe enormous variations between the other institutional measures we consider in our analysis.

The group of countries that export relatively fewer agricultural products demonstrates 1.77% (column 2) and 1.81% (column 4) of their GDP in resources rents. This aligns with the model of trade in Heckscher et al. (1991) that essentially states that countries export the products which use their relatively abundant and cheap factors of production. Hence, it is reasonable to expect that those economies less concentrated on the exports of commodities tend to exhibit a lower share of their GDP in natural resources

rents. Column 2 in table 1 also provides information indicating that a small portion (15%) of the countries in this group are landlocked yet most of them (85%) have access to the sea. The results in column 4 is pretty similar: 13% and 87%. Finally, the Gini index in table 1 discloses that countries with less focus on commodity exports have lower levels of inequality in respect of those whose exports are more intensive in agricultural varieties. Another interesting aspect we observe in table 1 is that countries that export fewer commodity goods show lower values of arable land, are less polluting, consume a larger amount of fertilizer, probably due to less favorable natural conditions, and spend more on research and development than those countries whose exports are more concentrated on agricultural goods.

In summary, we observe in table 1 that the group of countries with a higher intensity of commodity exports (both net exporters and those with a higher share of exports of that type of goods) demonstrate, on average, lower institutional quality levels, a higher percentage of natural resources rents, have access to the sea, present higher levels of inequality, have more arable land, pollute more, use relatively smaller amounts of fertilizers, and spend less on R&D. On the other hand, the group of economies with a lower intensity of commodity exports present, on average, higher institutional measures, a lower percentage of resources rents, can mostly have access to the sea, are less unequal, are endowed with smaller areas of arable land, emit less CO_2, consume more fertilizers, and show higher shares of R&D expenditures.

Table 1 also reports on the results of a test of equality of means to verify whether these two groups are statistically similar in the features we account for in this investigation. We note that the mean values in the group of countries with a higher intensity of commodity exports for almost all variables included in our analysis are statistically different from the control group (countries with lower intensity of agricultural exports) at 1%, 5%, or 10% of statistical significance. This emphasizes that both groups are somewhat heterogeneous and mostly dissimilar from one another.

As discussed in Imbens and Angrist (1994), this confirms that in this investigation treatment is not randomly assigned. That is, the decision of a country to become (or not), for example, a net exporter of agricultural goods is not random. Therefore, when we compare the background features of these countries we observe they are not similar

enough. In addition, we point out that in the context of this investigation, we are only able to unveil the local average treatment effect (LATE), that being the causal effect of commodity exports on the institutional development of those countries who responded to the Chinese demand shock for agricultural products. Windt (2023), for instance, claims that when the share of compliers is considerable and the treatment effects for different individuals (or countries, in our context) in the population are similar, then the LATE and the average treatment effect (ATE) do not differ widely. Therefore, since our groups are mostly, on average, statistically different in their background aspects, the results reported in table 1 confirm that this investigation discloses the LATE other than the ATE.

2.6.2. 2SLS regressions results

Based on the econometric model presented in equation (7), we obtain the empirical results and present them in this section. Notably, table 2 presents the first-stage regression results while table 5 reports the estimated coefficients of the second-stage regressions for each institutional dimension.

In this investigation, we look at the causal association between the exports of commodities and institutional quality. In our empirical specification, we are aware that the commodity products exports might, at the same time, cause and be affected by institutional quality. In addition, governmental strategies to boost the exports of agricultural goods and to improve institutional quality may be simultaneously determined by unobserved aspects, such as the motivation of a country leader, which would be included in the error term. Therefore, it is essential to search for exogenous variations in the exports of commodities in combination with econometric modeling to gain a comprehensive understanding of our causal relationship of interest.

We use the IV approach to deal with the endogeneity problem through the 2SLS estimator. The instrument in this study consists of the natural logarithm of the aggregate Chinese imports of agricultural goods. Since China represents one of the largest traders in the world (World Integrated Trade Solution, 2023), we believe that our instrument vastly determines the global foreign demand for agricultural products. Hence, this instrument captures the exogenous incentives that countries have to increase their exports of commodity goods; in other words, the Chinese imports of commodities affect our variable of interest.

In this context, the results of our first-stage regression, which involves regressing the endogenous variable (exports of commodities) on the instrumental variable and all remaining covariates, are presented in table 2. Afterward, the following paragraphs discuss the instrument strength, and explore the implications of its relevance.

Table 2 – First-stage regression results

Regressors:	Dependent variable: exports of commodities (USD)	
	(model 1)	(model 2)
Chinese imports of commodities (USD)	0.1584***	0.1372***
	(0.0273)	(0.0271)
Resources rent (% of GDP)	0.1741***	0.1712***
	(0.0130)	(0.0129)
Latitude (degrees)	0.0053***	0.0062***
	(0.0000)	(0.0013)
Landlocked (1 = yes; 0 = no)	-1.4256***	-1.3898***
	(0.0834)	(0.0860)
GDP *per capita* (USD)	1.2549***	1.3482***
	(0.0420)	(0.0483)
High-income country dummy		-0.2928***
		(0.0811)
Intercept	0.6505^{NS}	0.2814^{NS}
	(0.5299)	(0.5395)
Observations	1,225	1,225
R-sq	0.5525	0.5566
Adjusted R-sq	0.5507	0.5541
Partial R-sq	0.0266	0.0193
Robust F	33.6566	25.7289
Prob>F	0.0000	0.0000

Note: Robust standard errors in parentheses. ***p<0.001. NS non-significant. Exports of commodities, Chinese imports of commodities, and GDP *per capita* in log. Figures rounded to four decimal places.
Source: Own calculations.

Table 2 reports the results of the first-stage regression. Column 1 shows the estimations of our instrument as well as our control variables regressed on the exports of commodities (model 1). Column 2 includes a dummy variable to verify any possible changes in our estimations in case we account for higher-income economies[21] (model 2). Overall, we do not observe considerable variations after the inclusion of this binary

[21]We follow the country classification of the United Nations (2014) and create a dummy variable in our data to segregate developed from developing and transition economies. In our full sample, only North Macedonia and the Russian Federation are denoted transition economies. For simplicity purposes, we treat these two countries as developing nations.

covariate. Therefore, the analysis of the results in models 1 and 2 yield similar interpretations regarding the instrument relevance, which we discuss in the following paragraphs.

Our focus in table 2 is the estimated coefficient of the Chinese imports of commodities. This result confirms that the coefficients in models 1 and 2 for our instrument are statistically different from zero meaning that the Chinese international demand for agricultural goods affects the exports of commodity products. Moreover, the estimated coefficients show a positive association between the endogenous regressor and the instrumental variable, revealing that the Chinese demand shock increased the exports of commodities.

The results reported in table 2 work in favor of the relevance of the instrument (i.e., $Cov\left(chinaM_{i,t}, X_{i,t}\right) \neq 0$); however, some statistics derived from the first-stage regression provide an added indication of the instrument relevance, namely the R-sq, the Adjusted R-sq, the Partial R-sq, and the first-stage F statistic. The R-sq, for example, shows that the instrumental variable along with our controls determine a little more than 55% of the exports of commodities. Furthermore, for the corresponding first-stage regression with response 'exports of commodities', we might also want to know what percent of variation not explained by the Chinese imports of agricultural goods is explained by the control variables. The Partial R-sq indicates that the proportion of variation explained by our controls that cannot be rationalized by the instrument is low, 2.7% (model 1) and 2% (model 2). Additionally, a typical rule of thumb states that the F-stat for significance of the instrument in the first stage should exceed 10 (Stock et al., 2002). In table 2, the F test of excluded instrument is indeed more than 10 and reveals that we can reject the null hypothesis and confirm that all of the regression coefficients are not equal to zero.

Besides showing the results of the first-stage regression and providing some indication of the instrument relevance, it is also possible to conduct an underidentification test to verify whether the matrix rank is complete and, consequently, analyze the relevance of the excluded instrument. Table 3 reports these results.

Table 3 – Underidentification test

Kleibergen-Paap rk LM statistic (model 1):	32.2500
Chi-sq(1) P value:	0.0000

Kleibergen-Paap rk LM statistic (model 2):	24.9170
Chi-sq(1) P value:	0.0000

Source: Own calculations.

The results in table 3 confirm the relevance of the instrument used in this study. Specifically, we can observe that in both models 1 and 2 we are able to reject the null hypothesis that the matrix of reduced form coefficients has rank $= K - 1$. This means that the matrix has rank $= K$ and is identified, which corroborates the instrument relevance. However, we should interpret such a test with caution as it can still be the case that our models are only weakly identified and that our instrument is not strong. Thus, in table 4 we also perform a weak identification test in order to assess the instrument strength, if any.

Table 4 – Weak identification test (Cragg-Donald Wald F Statistic)

Stock-Yogo weak ID test critical values:	10%	15%	20%	25%
N = 1, K = 1	16.38	8.96	6.66	5.53
F-stat (model 1)				33.3380
F-stat (model 2)				23.9320

Note: The number of included endogenous regressors is represented by N, and the number of excluded instruments is represented by K.
Source: Own calculations.

The weak identification test results reported in table 4 confirm that our instrument is not weak. More clearly, the Cragg-Donald Wald F Statistic value shows no serious problem of estimated bias induced by weak IV. The F-test of excluded instruments is equal to 33.338 (model 1) and 23.932 (model 2) confirming that the instrument is strong.

In sum, the first-stage regression demonstrates that when we regress the exports of commodities on the Chinese international demand for agricultural products as well as on our control variables, we verify that the endogenous regressor is affected by the instrument in our empirical specification. Furthermore, we rely on some post estimation statistics, such as the R-sq, the Adjusted R-sq, the Partial R-sq, and the first-stage F-stat to validate the strength of our instrument. Lastly, we present the results of the identification and weak instrument tests, which confirm that that the Chinese demand shock serves as a relevant, strong instrument in our analysis.

Now we turn to our main results, the second-stage estimates of each institutional regression. Table 5 shows these results. We run different models alternating controls and the high-income country dummy. Specifically, column 1 shows the results for all institutional dimensions when we include no control variables or consider income structure. Column 2 exhibits estimates that account for other determinants of institutional quality with no control over the developing stage of countries. Finally, column 3, our preferred specification, yields results that not only considers control variables, but also the effect of being a more advanced economy. Additionally, columns 4 and 5 show regression results without and with control variables and correction for income structure through ordinary least squares (OLS), respectively. This allows us to have an idea of the direction of the resultant endogeneity bias that would prevail in the OLS parameter estimates for our endogenous variable. Coefficients for all control variables are provided in table 1A in the appendix.

Table 5 – IV regression results

	Endogenous regressor: exports of commodities (USD)				
Dependent variables:	2SLS			OLS	
	(1)	(2)	(3)	(4)	(5)
Voice and Accountability	-0.0282^{NS}	-0.7865***	-0.6054***		
	(0.1548)	(0.1548)	(0.1477)		
Political Stability and Absence of Violence	-0.4281^{NS}	-1.9081**	-1.5678*		
	(0.3723)	(0.9681)	(0.8962)		
Government Effectiveness	-0.1110^{NS}	-1.5514***	-1.5869***		
	(0.1385)	(0.4224)	(0.4399)		
Regulatory Quality	-0.0203^{NS}	-1.1449***	-1.1379***		
	(0.0467)	(0.2173)	(0.2478)		
Rule of Law	-0.0131^{NS}	-1.2420***	-1.1776***		
	(0.0470)	(0.2266)	(0.2468)		
Control of Corruption	0.0657^{NS}	-1.1700***	-1.2342***		
	(0.0806)	(0.2979)	(0.3263)		
Average institutional index	-0.0891^{NS}	-1.301***	-1.2183***	0.0909**	-0.1327*
	(0.0880)	(0.2898)	(0.2966)	(0.0425)	(0.0702)
High-income country dummy	No	No	Yes	No	Yes

Control variables	No	Yes	Yes	No	Yes
Observations	1,266	1,225	1,225	1,266	1,225

Notes: Robust standard errors in parentheses. ***p<0.01, **p<0.05, *p<0.1, and NS non-significant. Exports of commodities in log. Columns 1, 2 and 3 report the second-stage regression results without controls or the high-income country dummy, with control variables and no control over income structure, and with controls and high-income country dummy, respectively. Columns 4 and 5 report the OLS regression results without and with corrections for omitted variable bias, respectively, and no correction for endogeneity. Control variables are natural resources rents; log of per capita GDP; latitude, and the landlocked dummy. Figures rounded to four decimal places.

In table 5, the coefficients presented for all institutional dimensions in columns 2 and 3 are consistently statistically significant at the 1%, 5%, or 10% level. We find that increases in commodity exports do impact institutions negatively. We observe this outcome in all our estimations, regardless of the institutional dimension we consider. This result means that exporting agricultural goods causes institutional quality to suffer. Furthermore, when we combine the six measures of institutional quality into a single measure, the average institutional index, we also verify a negative causal association between commodity exports and institutions. Our results are somehow comparable to those in Lehne et al. (2014) and Islam and Montenegro (2002) who find that rule of law and control of corruption are often better in economies that do not have substantial natural resource endowments.

We emphasize that all the estimated coefficients in our study should be interpreted as the LATE commodities have on institutional quality. This means we unveil the impact of agricultural exports on the institutions of countries that are “moved” by the changes resulting from China’s accession to the WTO. While the ATE might seem ideal, the LATE provides a consistent estimate of the average treatment effect for a subgroup of the population: the compliers. The LATE is not able to capture the effects of the treatment, the Chinese demand shock, among all countries in our sample. However, this limitation is not necessarily a drawback; rather, the LATE could be exactly what we are interested in. That is, the average effect on those that actually comply with the assignment to the treatment. In our investigation, this is particularly motivating for we are mostly interested in revealing whether resource-rich economies, who tend to export commodities, respond to the Chinese demand shock for agricultural products, which in turn impacts their institutional development.

Trade theory gathers evidence that resource revenues can generate staggering

wealth that facilitates corruption and patronage networks and eliminate a critical link of accountability between citizens and government (Hodler, 2006; Karl, 1997). Moreover, according to the Heckscher-Ohlin model, it is also known that countries tend to specialize in the commerce of those goods that use their abundant production factor more intensively. For this reason, our results are in line with the related theory in that they unveil a causal association suggesting that commodity international trade (as a consequence of natural resources abundance) negatively impacts institutional development. Moreover, this paper discloses that the natural endowment – institutions nexus extends to the exports of agricultural products and institutional quality.

Turning to the magnitude of the causal effect of commodity exports onto the quality of institutions, we concentrate our interpretation on column 3, our preferred specification. These results not only account for other determinants of institutional quality, but also the income structure of the sample. We find that government effectiveness and political stability more affected by variations in the exports of commodity goods than other institutional dimensions, such as voice and accountability. More clearly, increases in agricultural exports impacts political stability almost two times more than voice and accountability. Again, Karl (1997) observes that ruling groups may wield influence on politics in favor of natural resource rents hampering institutional development. Similarly, it is reasonable to consider that exploiting natural resources aiming at higher volumes of commodity production and its exports would then have a lasting negative effect onto institutional quality.

Our results are consistent to different sensitivity checks mainly through columns 1 and 3. Specifically, estimates in column 3 in table 5 show that percentage increases in agricultural exports distress institutional strength indexes in at least 0.61. When we rely on the average institutional index, we lose some clarity as to which dimensions are more severely impacted as our sampled economies export higher volumes of commodity goods. However, these estimates reported in column 3 reassure the negative causal link running from commodity exports to the quality of institutions.

In table 5, column 3 includes control variables and the high-income country dummy to try and eliminate omitted variables bias. One common approach to investigating mediators of an economic relationship, however, is to also run regression models with

only the variable of interest predicting the outcome and then have an indication of the direction of such a bias. For this reason, columns 1 and 2 exhibit estimation results without any control variables. Note that in column 1, for instance, all estimated coefficients are statistically not different from zero while the same parameter estimates in our preferred specification (column 3) exhibit highly significant figures. This change in the coefficients of commodity exports represents the result of the omitted variables bias. By including controls, we are able to identify in table 4 the proportion of the total effect of commodity trade that is explained by our control variables and the proportion that remains solely attributable to commodity exports.

Lastly, columns 4 and 5 display parameter estimates of the average institutional index regressed on our variable of interest without and with controls and income structure dummy through OLS, respectively, leading us to biased coefficients. We can verify that had we not considered such econometric issues, our evaluation would lead us to a positive effect (column 4) and an underestimate (column 5) of the direct effect of commodity trade onto institutions. The OLS (biased) estimate, reported in 5, for example, suggests that a 1% increase in commodity foreign trade negatively affects the average institutional quality index by only 0.13. However, our IV regression results, shown mainly in column 3, clarify that such an effect is in fact much bigger, i.e. commodity exports cause a reduction in the average institutional quality of 1.22. Therefore, our OLS outcomes in columns 4 and 5 provide us with the direction of the resultant endogeneity bias while our identification strategy allows us to deliver the causal relationship between commodity exports and institutional quality.

In summary, the results we present in table 5 suggest that as countries export larger volumes of commodities, or as countries focus on a commodity based exporting product bundle, they are negatively affecting their citizen's participation in selecting their government (voice and accountability) as well as promoting the likelihood of politically motived violence (political stability and absence of violence). Moreover, the results in table 4 unveil that our variable of interest has a negative impact on the quality of public services (government effectiveness) and on the ability of the government to formulate and implement sound policies that permit and promote private sector development (regulatory quality). Last but not least, we find that the trade of commodity products undermines the

confidence in the rules of society (rule of law) and increase the extent to which public power is exercised for private gain (control of corruption)[22].

Overall, this study corroborates, at least to some extent, existing findings that countries with abundant natural resources often experience the worst development outcomes, characterized by poverty, inequality, and deprivation (Mehlum et al., 2006; Robinson et al., 2006), that could be a consequence of lagging institutional maturity. Lehne et al. (2014) observe a negative impact of natural resource abundance relative to a country's size on institutional development, which we could verify in the relationship between the exports of resource intensive products and institutions. Moreover, our results align with Islam and Montenegro (2002), who find that economies lacking substantial natural resource endowments tend to have better rule of law and lower levels of corruption. Hodler (2006) provides further context, explaining that natural resources can lead to conflicts among competing groups, reducing production and weakening property rights, thereby diminishing the attractiveness of productive activities. All in all, these studies reveal a negative impact of natural resource endowment on institutional development, and since resource-rich countries often export goods reliant on these resources, their findings are consistent with our research.

2.6.3. 2SLS regressions results: Disaggregated exports

The results we present in this section go one step further in the understanding of the causal association between agricultural exports and institutional quality. We are aware that there exist varying types of commodities that could impact institutions differently. Therefore, we also verify whether increases in the exports of animals, vegetables, food products, minerals, and fuels influence institutional quality similarly to the aggregate causal effect we unveil in table 5. The selection of these groups represents the standardized product bundles in WITS (2023), namely, animals (all products ranging from HS01 and HS05); vegetables (HS06-HS15); foods (HS16-HS24); minerals (HS25-HS26); and fuels (HS27). For these estimation, we adjust our instrument and consider Chinese imports of these same product classifications. Table 6 shows the results for the first-stage relationship and provides some indication of the instrument relevance. We also run an

[22] Documentation on each institutional dimension is available in the WGI database. See https://info.worldbank.org/governance/wgi/Home/Documents#wgiDataCrossCtry

underidentification test to verify whether the matrix rank is complete and, consequently, analyze the relevance of the excluded instrument.

Table 6 – First-stage results: disaggregated exports

Instrument: Chinese imports (USD)	Dependent variables: exports (USD)				
	Animals	Vegetables	Foods	Minerals	Fuels
Animals	0.1563***				
	(0.0365)				
Vegetables		0.3344***			
		(0.0419)			
Foods			0.1925***		
			(0.0321)		
Minerals				0.1388***	
				(0.0373)	
Fuels					0.1272***
					(0.0489)
Intercept	0.0596^{NS}	2.2397***	0.5132^{NS}	-0.0791^{NS}	-5.3892***
	(0.6469)	(0.7543)	(0.5704)	(0.8327)	(0.8617)
Observations	1,224	1,225	1,225	1,225	1,225
R-sq	0.4414	0.2466	0.4585	0.245	0.416
Adjusted R-sq	0.4387	0.2429	0.4558	0.2412	0.4131
Partial R-sq	0.0142	0.0473	0.0263	0.0100	0.0049
Robust F	18.3750	63.6584	36.0684	13.8269	6.7774
Prob>F	0.0000	0.0000	0.0000	0.0000	0.0093
Kleibergen-Paap rk LM stat	17.7620	59.2360	33.8900	13.5330	6.7580
Chi-sq(1) P value	0.0000	0.0000	0.0000	0.0002	0.0093
Stock-Yogo weak ID test (Cragg-Donald Wald F stat)					
F-stat	17.4870	63.6580	32.8550	12.2570	5.9470

Notes: Robust standard errors in parentheses. ***p<0.01, **p<0.05, *p<0.1, and NS non-significant. Exports and Chinese imports in log. Control variables included: natural resources rents; log of per capita GDP; latitude, and the landlocked dummy. Figures rounded to four decimal places. The number of included endogenous regressors is represented by N=1, and the number of excluded instruments is represented by K=1.
Source: Own calculations.

The results we present in table 6 similarly to those reported in table 5 indicate that there exists a strong and statistically significant relationship running from the exports of animals, vegetables, food products, minerals, and fuels to the Chinese imports of these product bundles. More clearly, the estimated coefficients evidence a positive association

between the endogenous regressors and our instruments, which disclosed that the Chinese demand shock had a positive effect on commodity exports. The coefficients in table 6 unveils that such an effect is more prominent on the exports of vegetables, such as live trees, roots, fruits, coffee, etc., and food products, like preparations of meat, sugars, cocoa, and tobacco.

The estimated coefficients in table 6 favor the relevance of our instrument; i.e.: the covariance between the Chinese imports of each product bundle and the exports of disaggregated commodities is different than zero. In addition, table 6 provides some statistics from the first-stage relationship that support the relevance of the instrument, such as the R-sq, the Adjusted R-sq, the Partial R-sq, and the first-stage F statistic. For instance, the R-squared value illustrates that both our instrumental variable and control factors collectively account for at least 24.5% of the variance in commodity export. also observe that the R-squared is much higher when the output variable corresponds to the exports of food and animal products.

The Partial R-squared values helps to identify what portion of the unexplained variation in Chinese agricultural goods imports can be attributed to our control variables. The Partial R-squared values shed light on this, indicating a low proportion of 4.7% (exports of vegetables), for example, for the variance explained by our controls but not accounted for by the instrument. Moreover, a common rule of thumb suggests that the F-statistic for instrument significance in the first stage should surpass 10 (Stock et al., 2002). In table 5, the F test for the excluded instrument indeed exceeds this threshold for all regressions, except for the exports of fuels, allowing us to reject the null hypothesis and affirm that none of the regression coefficients are equal to zero to almost all our estimations. Therefore, we note that our instrument has little significance when we consider the exports of fuels, meaning that we cannot confirm that the Chinese demand for this product bundle strongly correlates with the exports of fuels originated in the sampled countries.

In table 6, besides reporting the results of the first-stage regression and providing some indication of the instrument relevance, we also present the outcome of an underidentification test to verify whether the matrix rank is complete and evaluate the relevance of the excluded instrument as well as a test to verify whether our instrument is

weakly identified or not. The results in table 6 mostly confirm the relevance of the instrument used in this analysis. Precisely, we observe that with the exception of fuels exports, we are able to reject the null hypothesis that the matrix of reduced form coefficients has rank $= K - 1$. This means that the matrix has rank $= K$ and is identified, which substantiates the instrument relevance. Finally, the weak identification test results ratify that our instrument is mostly adequate. More clearly, the Cragg-Donald Wald F Statistic value shows no serious problem of estimated bias induced by weak IV mainly for the regressions considering the exports of animal, vegetable, and food products. The F-test of excluded instruments is equal to 17.4870 and 63.6580, and 32.8550 confirming that the instrument is strong. For the remaining regressions, namely, exports of minerals and fuels the F-test of excluded instruments yields 12.2570 and 5.9470 meaning that these estimates could be affected by biases to the order of 15% and 25%, respectively, according to Stock-Yogo weak ID test critical values (see table 4).

Now we delve into the second-stage estimates that reveal the causal effect running from different commodity product bundles to the average institutional quality index. These results are showcased in table 7. For simplicity, we only report coefficients for our variables of interest, namely, exports of animals, vegetables, foods, minerals, and fuels. However, all regressions include covariates to control for other observable determinants of institutional quality as well as the high-income country dummy.

Table 7 – IV regression results: Disaggregated exports

Endogenous regressor: exports (USD)	2SLS				
	Dependent variable: average institutional quality				
Animals	-2.1680**				
	(0.8427)				
Vegetables		-0.6561***			
		(0.1211)			
Foods			-1.1506***		
			(0.2463)		
Minerals				-1.3171***	
				(0.4506)	
Fuels					-2.4907^{NS}
					(1.5905)
Intercept	-2.5420^{NS}	-0.7736^{NS}	-1.7862^{NS}	-3.4636***	-16.0499**

	(1.6186)	(1.2227)	(1.1659)	(1.2789)	(7.4661)
High-income country dummy	Yes	Yes	Yes	Yes	Yes
Control variables	Yes	Yes	Yes	Yes	Yes
Observations	1,224	1,225	1,225	1,225	1,225
Wald chi2(5)	41.52	511.2	275.57	66.28	17.34
Prob > chi2	0.000	0.000	0.000	0.000	0.000

Notes: Robust standard errors in parentheses. ***p<0.01, **p<0.05, *p<0.1, and NS non-significant. Exports in log. Control variables are natural resources rents; log of per capita GDP; latitude, and the landlocked dummy. Figures rounded to four decimal places.

Table 7 demonstrates that the estimated coefficients with the exception of fuels exports are statistically significant at 1% or 5% level of statistical significance. Our findings consistently indicate that an increase in the exports of animal, vegetable, food, and minerals has a detrimental impact on the average institutional index. This observation holds true across nearly all our estimations, non-significant for the exports of fuels. On average, we verify that exporting animal products negatively impacts institutions more strongly vis-à-vis the types of commodity exports. Essentially, exporting commodity goods appears to lead to a decline in institutional quality, with a greater impact when countries export animal varieties and less significant impact when the sampled economies ship vegetable products.

We comment on the rationale of such result. Raising cattle, for instance, involves a significant consumption of natural resources, ranging from land and water to feed and energy. The scale of resource utilization in this industry is substantial, making it a resource-intensive endeavor. However, the ability to engage in cattle farming, especially on a large scale, often hinges on the influence of powerful producer groups or individuals who have the capacity to influence politicians and shape legislation to their advantage. This influence can lead to the more flexible laws and regulations that favor the interests of these producer groups. Therefore, we highlight that our results could be attributed to this behavior; that is, the impact of such lobbying efforts on institutional quality. When producer groups successfully advocate for more permissive regulations, it can lead to a weakening of institutional checks and balances. This can manifest in various ways, including reduced environmental safeguards, relaxed labor standards, or decreased oversight and accountability, which could altogether be reflected onto our outcome variable in table 7, the average institutional index. As a result, the overall quality of

institutions can be compromised, potentially leading to negative consequences for society at large.

The disparities in the impact of lobbying by producer groups on institutional quality between animal (-2.168) and vegetable exports (-0.6561), for example, could stem from several key factors. Cattle farming, being highly resource-intensive and often associated with contentious issues like environmental degradation and animal welfare, encourages powerful producer groups to exert substantial influence on politicians and legislation. In contrast, the vegetable industry is characterized by a relative lower resource intensity, a more diverse array of products and stakeholders, and a generally less controversial public image. These factors collectively could lead to a smaller lobbying effect on institutional quality in the vegetable sector, providing some insights towards the findings in table 7.

The lack of a significant effect of fuel exports, such as oil or natural gas, on institutional quality in table 7 may be attributed to a range of factors. Countries reliant on fuel exports may implement strategies to mitigate the resource curse, promote economic diversification, and maintain political stability, diminishing the negative impact on institutions. Additionally, external factors like global oil prices and government policies allocating resource revenues can play pivotal roles in shaping institutional outcomes. The influence and lobbying tactics of producer groups in the fuel sector may differ from other industries, and there may be a time lag in observing the effects on institutional quality.

In a nutshell, table 7 presents statistically significant coefficients for various export product categories, except for fuel exports, indicating a consistent damaging impact on average institutional quality from increased exports of animal, vegetable, food, and mineral goods. The disproportionate impact of animal exports compared to vegetable exports could be explained by the resource-intensive nature of cattle farming, which relies on relatively more powerful producer groups that could influence legislation in their favor, potentially leading to weakened institutional checks and environmental safeguards. Conversely, the vegetable industry's rather lower resource intensity, diverse products, and less controversial image may result in a smaller lobbying effect on institutional quality. The insignificant effect for fuel exports may be due to countries implementing strategies to mitigate the resource curse, promote diversification, and maintain stability, along with external factors like global oil prices and different lobbying dynamics in the fuel sector.

2.7. Concluding remarks

Along with the development of the global economy and the constantly changing trade dynamics, the incentives towards the production of agricultural goods may be a decision made by various governments worldwide. In this context, this article sheds light on the causal relationship between commodity exports and institutional quality in a set of 49 countries spanning from 1997 to 2022.

The main findings underscore the far-reaching impact of variations in agricultural exports on a country's institutional landscape. Specifically, as nations increasingly engage in commodity exports or rely on commodity-based product bundles as a response to the Chinese shock, they tend to diminish their citizens' participation in government selection (voice and accountability) and elevate the risk of politically motivated violence (political stability and absence of violence). Moreover, our results reveal a disadvantageous effect on public services quality (government effectiveness) and the government's ability to formulate and implement policies promoting private sector development (regulatory quality). Notably, the trade of commodity products erodes confidence in societal rules (rule of law) and amplifies the extent to which public power is exploited for private gain (control of corruption).

Earlier studies show that there exists a negative relationship between commodity wealth and institutional development. In this paper, we confirm that the natural resource – weak institutions association also holds in the context of commodity exports and institutional quality. Second, we unpretentiously add to the debate over the resource-curse hypothesis by disclosing that commodity trade exporters, who tend to be rich in natural resources, indeed impact their institutional levels negatively. Third, by using a natural experiment, we are capable of revealing a causal relationship. That is, our results consistently indicate that as countries focus upon commodity based exporting goods, they cause a reduction in the effectiveness of their institutions. Fourth, we use varying appropriate measures of institutional quality, enhancing the quality of our results.

The implications of our findings for economic policy are substantial. Firstly, policymakers in commodity-exporting countries should be conscious of the potential trade-offs between economic gains from exports and the quality of their institutions. While resource exports can boost revenues, they may also lead to institutional deterioration,

affecting governance, public services, and political stability. Thus, governments might consider the adoption of measures to mitigate these adverse effects, such as transparent revenue management, diversification of their economies, and robust regulatory frameworks. Secondly, our research underscores the need for differentiated policies based on the type of commodities being exported. Cattle farming, for instance, requires special attention due to its resource-intensive nature and the lobbying power of producer groups. Conversely, vegetable exports may warrant more relaxed regulation. Lastly, our results highlight the importance of external factors, like global oil prices, in shaping institutional outcomes. Policymakers should be attuned to these factors and develop strategies that safeguard institutions in the face of volatile global commodity markets.

2.8. Appendix

Table 1A – Estimation coefficients for IV regression controls

	VAE (2SLS)		PVE (2SLS)		GEE (2SLS)		RQE (2SLS)		RLE (2SLS)		CCE (2SLS)		AVG (2SLS)		AVG (OLS)	
Commodity exports (USD)	-0.787***	-0.605***	-1.908**	-1.568*	-1.551***	-1.587***	-1.145***	-1.138***	-1.242***	-1.178***	-1.170***	-1.234***	-1.301***	-1.218***	-0.158***	-0.132*
	(0.155)	(0.148)	(0.968)	(0.896)	(0.422)	(0.434)	(0.217)	(0.248)	(0.227)	(0.247)	(0.298)	(0.326)	(0.290)	(0.297)	(0.075)	(0.070)
Resources rent (% of GDP)	0.096***	0.068**	0.316*	0.265^{NS}	0.269***	0.275***	0.181***	0.180***	0.196***	0.186***	0.190***	0.200***	0.208***	0.196***	-0.002^{NS}	0.002^{NS}
	(0.030)	(0.027)	(0.181)	(0.169)	(0.080)	(0.082)	(0.041)	(0.045)	(0.043)	(0.046)	(0.056)	(0.060)	(0.055)	(0.055)	(0.015)	(0.015)
Latitude (degrees)	0.0004^{NS}	-0.002^{NS}	0.019^{NS}	0.014^{NS}	0.010^{NS}	0.010*	0.004**	0.004*	0.004^{NS}	0.003^{NS}	-0.001^{NS}	-0.001^{NS}	0.006^{NS}	0.005^{NS}	0.0004^{NS}	-0.002^{NS}
	(0.002)	(0.001)	(0.018)	(0.017)	(0.006)	(0.006)	(0.002)	(0.002)	(0.002)	(0.002)	(0.004)	(0.004)	(0.004)	(0.004)	(0.181)	(0.003)
Landlocked (1 = yes; 0 = no)	-1.013***	-0.803***	-2.606^{NS}	-2.212^{NS}	-2.229***	-2.270***	-1.542***	-1.534***	-1.611***	-1.536***	-1.508***	-1.582***	-1.751***	-1.656***	-0.139^{NS}	-0.172^{NS}
	(0.229)	(0.208)	(1.685)	(1.595)	(0.694)	(0.707)	(0.325)	(0.357)	(0.339)	(0.354)	(0.475)	(0.503)	(0.466)	(0.464)	(0.181)	(0.184)
GDP *per capita* (USD)	1.469***	1.116***	2.821**	2.157**	2.771***	2.840***	2.187***	2.173***	2.379***	2.253***	2.388***	2.513***	2.336***	2.176***	0.792***	0.592***
	(0.215)	(0.221)	(1.099)	(0.976)	(0.523)	(0.573)	(0.297)	(0.360)	(0.311)	(0.360)	(0.383)	(0.445)	(0.368)	(0.397)	(0.054)	(0.031)
High income country dummy		0.396***		0.744***		-0.077^{NS}		0.015^{NS}		0.141^{NS}		-0.140^{NS}		0.180^{NS}		0.602***
		(0.083)		(0.248)		(0.177)		(0.121)		(0.123)		(0.138)		(0.133)		(0.090)
Intercept	-1.069**	-0.688*	3.070^{NS}	3.786^{NS}	-1.761^{NS}	-1.836^{NS}	-2.401***	-2.386***	-2.854***	-2.719***	-4.105***	-4.240***	-1.520^{NS}	-1.347^{NS}	-4.508***	-3.275***
	(0.524)	(0.415)	(4.744)	(4.879)	(1.842)	(1.872)	(0.741)	(0.699)	(0.810)	(0.751)	(1.203)	(1.227)	(1.213)	(1.194)	(0.711)	(0.875)
Observations	1,225	1,225	1,225	1,225	1,225	1,225	1,225	1,225	1,225	1,225	1,225	1,225	1,225	1,225	1,225	1,225
Wald chi2	224.47	556.09	72.81	166.89	155.78	154.57	266.38	278.90	239.22	296.97	260.45	257.82	188.63	252.33	445.49	508.22
Prob>chi2	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000

Notes: Robust standard errors in parentheses. ***p<0.01, **p<0.05, *p<0.1, and NS non-significant. Exports of commodities in log. Figures rounded to three decimal places.

3. DOES ECONOMIC POLICY UNCERTAINTY AFFECT COMMODITY-STOCK MARKET INTERACTIONS IN BRAZIL?[23]

Victor Henrique Lana Pinto, Leonardo Bornacki de Mattos[24]

Abstract
At times of economic volatility and policy unpredictability, understanding the nuanced relationship between uncertainty and the correlations observed within commodity and stock markets becomes imperative. Drawing on a comprehensive review of existing literature and employing econometric techniques, we investigate the dynamic interactions between commodity futures (live cattle and soybeans) and the Brazilian stock market, and whether they respond to heightened uncertainty originating in Brazil and abroad. First, we rely on a DCC-GARCH model to estimate the time-varying correlations between live cattle-, soybeans-, and stocks from January 2014 to July 2023. Then, we evaluate how economic policy uncertainty influences these correlations through the ordinary least squares estimator with Newey-West standard errors correction. Additionally, we examine the joint effect of uncertainty and economic conditions (industrial production growth, economic activity index, interest rate, and exchange rate) in Brazil on the outcome variable. We detect a statistically positive yet negligible effect of domestic and international economic uncertainty on the dynamic interactions between the commodity and stock markets in Brazil. Moreover, we largely verify that economic conditions in Brazil tend to intensify the time-varying correlations between live cattle and stocks relatively more than the interactions between soybeans futures and the Brazilian stock market. We conclude that the low liquidity of commodity futures on B3, especially soybeans, might lead investors to rely on alternative market fundamentals to mitigate uncertainty, reducing its impact on the interactions between Brazil's commodity and stock markets.

Keywords: Economic Policy Uncertainty; Live Cattle; Soybeans; Ibovespa; DCC-GARCH

JEL codes: C32; Q02; Q13

3.1. Introduction

The relevance of researches evaluating whether economic policy uncertainty may affect financial markets is noteworthy. Unforeseen changes might delay hiring and investments, causing a reduction in economic growth, a decline in profitability, and a decrease in asset prices (Smales, 2021), especially in developing economies where uncertainty is more recurrent (Ahir et al., 2022). In Brazil, for instance, corruption scandals, a presidential impeachment, and elections with competing parties from opposite ends of

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the political spectrum have exposed the country to economic uncertainties (Ferreira et al., 2019). Furthermore, a broad range of world frictions such as financial (e.g. global crisis), and health (e.g. Covid-19) related events could influence investment decisions allowing for considerable policy uncertainty. This means that both domestic and/or global instability play a major role in how investors diversify their asset classes (Smales, 2021).

Among many portfolio strategies, investors have begun considering agricultural goods into their investment plans combining commodities and stocks to achieve diversified returns (Wen et al., 2021) . Because commodity price drivers such as world demand, productivity growth rates, climate, geopolitics, and supply constraints are different from those that determine the value of stocks, the interactions between the commodity and stock markets assume greater significance to portfolio diversification returns (Daskalaki & Skiadopoulos, 2011; Gorton & Rouwenhorst, 2006).

The associations between commodity futures contracts and the stock market can reveal important insights into the connections of financial and agricultural markets. Positive correlations between commodities and stocks often reflect a period where the market sees them as influenced by similar factors, while a negative correlation between commodities and stocks is typically viewed optimistically by investors seeking to reduce risk and add stability to their portfolios. Understanding this relationship is useful because it can help investors gauge the potential impact of stock market volatility on commodity prices and vice versa. Moreover, these correlations may reflect how economic shocks influence both markets, providing valuable information for risk management, portfolio allocation, and forecasting future market movements. Therefore, studies dedicated to understanding whether the interactions between the commodity and stock markets vary in response to economic policy uncertainty become increasingly relevant, mainly in Brazil, a country that has experienced various instability episodes in the past decades (Ferreira et al., 2019).

Various researchers have investigated whether financial markets respond to uncertainty yet others have explored cross-market spillover effects from the interactions among different market assets. Arouri et al. (2016), Bekiros et al. (2016), and Smales (2021) provide ample evidence for a policy uncertainty effect on the stock market while Antonakakis et al. (2014), Wang et al. (2015), and Watugala (2019) evaluate how

economic uncertainty affects commodity markets. Geng et al. (2021) and Zhang (2017) focus on the dynamic relationship between the stock and energy markets, while Choi and Hammoudeh (2010), and Creti et al. (2013) analyze volatility spillovers in the commodity and stock markets. In sum, these studies observe the existence of an isolated effect running from economic instability episodes towards commodities and the stock market, and also, the stock-energy interactions as well as the correlations between agricultural and stock markets. Nevertheless, there is still room for investigations aiming to understand if economic policy uncertainty could influence the interactions between different market assets, especially for a developing, more vulnerable economy such as Brazil.

Some studies assess how the correlations among various financial asset classes respond to economic policy uncertainty. For instance, Fang et al. (2017) and Li et al. (2015) explore the policy uncertainty effect on the US stock and bond markets connections. Both studies document that innovations in the policy uncertainty index has a negative influence on the US stock-bond markets correlations. Alternatively, Fang et al. (2018) demonstrate how policy changes affect the correlations between the stock and oil markets and find a positive effect on the long-run interactions between those two markets. Similarly, Badshah et al. (2019) examine whether economic policy instability in the US affects the dynamic conditional interactions between the commodity and stock returns. The authors discover that economic policy uncertainty is positively associated with commodity-stock correlations and that this influence is particularly strong for energy and metal commodities. Lastly and more recently, Zhao and Wang (2022) evaluate the effects of economic and monetary policy uncertainty in the US and China on oil-stock and gold-stock correlations. They find that gold can provide a better diversification for stock market risks than crude oil during higher levels of economic uncertainty.

Within the context of Brazil, the live cattle and soybeans markets stand out as key components of the nation's economy, reflecting its agricultural prowess and global influence. Brazil's status as a major exporter of beef and soybeans underscores the importance of these commodities, potentially reflected in the trading of their respective futures contracts on the Brazilian stock exchange, the B3. This, in turn, could potentially act as instruments for price discovery and risk management for producers and traders alike. From 2014 to 2023, for instance, the average daily price of a live cattle futures

contract on B3 hovered around USD 39, with a daily traded volume of 1,281 contracts. In addition, soybeans futures contracts averaged nearly USD 24 daily, with a traded volume of 209 contracts[25].

The dynamics of these markets frequently intersect with Brazil's stock market, reflecting broader economic sentiment and investor confidence. However, this interaction also makes them vulnerable to economic uncertainty, not only within Brazil but also from major economies like the US and China. Fluctuations in commodity prices can ripple through financial markets, influencing investor behavior and market stability. As a result, understanding the complex relationship between these futures markets and the broader economy is crucial for navigating the intricacies of Brazil's financial landscape.

In addition to the potential effects of uncertainty on cross-asset markets, Pástor and Veronesi (2013) highlight that political shocks wield greater influence over asset prices when the economy is in a more fragile state. This implies that as economic conditions deteriorate, uncertainties surrounding policy and governance can have a heightened impact on market dynamics. Therefore, it becomes imperative to evaluate how fluctuations in uncertainty might affect the interconnectedness of commodity and stock markets as economic drivers such as interest and exchange rates, production growth, and overall economic activity alter.

In this paper, we look to extend the related literatures on the association between economic uncertainty and commodity-stock correlations by attempting to answer the following questions: First, does domestic economic policy uncertainty have substantial effects on time-varying correlations between live cattle-, soybeans- and the Brazilian stock market? Second, does economic uncertainty in the US, China, and the world affect how these commodity futures contracts interact with the stock market in Brazil? Third, is the effect of uncertainty intensified during particular economic conditions? Therefore, this paper aims to examine whether the dynamic correlations between commodity futures (live cattle and soybeans) and the stock market respond to economic policy shocks originated both in Brazil and abroad from January 2014 to July 2023. This timeframe encompasses

[25] The agricultural futures contracts traded on B3 include arabica coffee, live cattle, corn, and soybeans. From 2014 to 2023 the average daily closing prices for these commodities on B3 were USD176.25 (coffee), USD39 (live cattle), USD24 (soybeans), and USD10 (corn). Live cattle and corn futures contracts on B3 are traded in Brazilian reals, which, for the purpose of broad comparison, were roughly converted into US dollars using an exchange rate of BRL5. During the same timespan, the average traded volume reached 3,593 (corn), 1,281 (live cattle), 367 (coffee), and 209 (soybeans) contracts.

both domestic and international events that influenced Brazil's economic environment, including corruption scandals and the Covid-19 pandemic, while also considering the data available at the time of the econometric analysis. In addition, we look to verify the existence of a joint effect of policy uncertainty and certain economic states on cross-market, commodity-stock correlations.

This paper contributes to the literature by exploring the impact of economic policy uncertainty on financial market correlations within a developing economy context, specifically Brazil – an area that remains underexplored compared to studies in advanced economies. It offers an empirical analysis of the correlations between two major Brazilian commodities (live cattle and soybeans) and the stock market (Bovespa Index[26]), assessing their potential as hedges or diversifiers against market volatility. The study further investigates the spillover effects of foreign uncertainty on Brazil and how varying economic conditions influence these correlations.

Overall, our results reveal a positive yet marginal influence of domestic and foreign economic policy uncertainty on the interactions between live cattle-, soybeans- futures and the Brazilian stock market. We attribute this outcome to specific characteristics of Brazil's commodity futures market, particularly its relatively low volume and liquidity. It is likely that uncertainty, both domestic and international, does not significantly impact commodity-stock interactions in the country because investors may not heavily rely on commodity futures as a primary means of securing their investments against stock market volatily.

In that sense, our findings offer insights with implications for policy. Policymakers may recognize that while uncertainty might not strongly influence commodity-stock interdependence in Brazil, other factors could be more significant in shaping this relationship. Investors can still benefit from these insights by focusing on assets with limited spillover effects, enhancing their risk management strategies. Thus, this study aids Brazilian corporate managers and decision-makers in understanding the dynamic between commodity and stock markets, even in contexts where economic policy uncertainty has a subdued role in driving correlations.

[26] The Bovespa Index or Ibovespa is the main performance indicator of the stocks traded on B3, one of the largest exchanges in the world in terms of market value, the second largest in the Americas, and the leading exchange in Latin America. Ibovespa lists major companies in the Brazilian capital market (B3, 2022).

The remainder of this paper is structured as follows: Section 3.2 provides a concise theoretical framework. Section 3.3 details the dataset utilized for analysis. Section 3.4 outlines our research design and methods. Section 3.5 presents the empirical findings and subsequent discussion. Finally, Section 3.6 offers concluding remarks of the paper and section 3.7 presents the appendix.

3.2. Theoretical framework

This section offers the conceptual framework for the effect of the economic policy uncertainty on the correlations between the commodity and stock markets. First, it briefly presents the theory set out in Pástor and Veronesi (2013) for the association between uncertainty and financial markets. After that, based on the portfolio selection model of Markowitz (1952) and on the related literature, it provides theoretical insights into the links between the commodity and stock markets.

The model of Pástor and Veronesi (2013) considers an economy with a finite horizon $[0,T]$ and a range of firms $i \in [0,1]$. The capital of firm *i* at time *t* is B_t^i. Because firms are entirely financed by equity, B_t^i can be perceived as book value of equity. When $t = 0$, all firms place an equal amount of capital, which Pástor and Veronesi (2013) normalize to $B_0^i = 1$. The capital of each firm *i* is invested in a linear technology and its stochastic rate of return is $d\Pi_t^i$. Also, all profits are reinvested in a way that the capital of each firm *i* grows according to $dB_t^i = B_t^i d\Pi_t^i$. The profitability of the firms is then given by $d\Pi_t^i$. Formally, for all $t \in [0,T]$, profitability follows the process:

$$d\Pi_t^i = (\mu + g_t)d_t + \sigma dZ_t + \sigma_1 d_t Z_t^i \quad (1)$$

where μ, σ, and σ_1 are observable constants, Z_t refers to a Brownian motion[27], and Z_t^i also indicates a Brownian motion that is specific to firm *i.* The variable g_t represents the impact of the current governmental policy on the average profitability process of each firm *i.* Pástor and Veronesi (2013) interpret these policies as government actions that affect the economic environment in which the private sector operates. For the authors, changes in these policies impact firms by levying taxes, enforcing laws, recommending subsidies,

[27] Brownian motion denotes a significant concept in economics in two respects. It underlies an important part of stochastic finance, which includes the pricing of risky assets, such as stock prices, bonds and exchange rates. Besides, it plays a central role in econometrics, especially in the distribution theory underlying unit root test statistics (Patterson, 2010).

implementing environmental policies, regulation competition, etc. Governments, in turn, change these actions from time to time, causing uncertainty and price reactions in financial markets. In the context of this study, we could interpret g_t as the economic policy uncertainty and the profitability of the firms as the returns of the financial market, or how commodity futures interact with stock prices. When $g_t = 0$, the economic policy is considered neutral and does not affect profitability.

The impact of the economic policy g_t is constant while the same policy is in effect. At a random exogenous given time τ, $0 < \tau < T$, the government may decide to change the policy. Therefore, g_t is a function of time, accordingly:

$$g_t = \begin{cases} g^{old} \; for \; t \leq \tau & \\ g^{old} \; for \; t > \tau & \text{if there is no policy change} \\ g^{new} \; for \; t > \tau & \text{if there is policy change} \end{cases} \quad (2)$$

In (2), g^{old} indicates the effect of a policy prevailing at time 0. A policy change substitutes g^{old} for g^{new}, inducing a shift in the average profitability of firms. Pástor and Veronesi (2013) assume that the values of g^{old} and g^{new} are unknown. This is a fundamental assumption that grasps the idea relating uncertainty to firm profitability. The distributions of g^{old} and g^{new} when $t = 0$ are assumed to be conditional normally distributed with mean zero and variance σ_g^2; that is, $g \sim N(0, \sigma_g^2)$. The authors refer to σ_g as the 'impact uncertainty'. Lastly, the value of g_t is not known for all $t \in [0, T]$ to the government and to investors who own the firms.

Turning to the firms, they are owned by a continuum of investors who seek to maximize expected utility. This function for each investor *j*, where $j \in [0,1]$, follows:

$$u\left(W_T^j\right) = \frac{(W_T^j)^{1-\gamma}}{1-\gamma} \quad (3)$$

where W_T^j represents the wealth of investor *j* in T and $\gamma > 1$ indicates the risk aversion coefficient. When $t = 0$, every investor possesses the same number of shares of firm *i*. These shares pay dividends in T and investors watch whether a policy may occur in τ.

The government adopts a policy at time τ maximizing the same objective function as investors. However, policymakers need to embed a cost (or benefit) related to the policy change. More clearly, the government introduces a policy when the expected utility under the new policy is higher than the old policy. Formally:

$$\max\left\{E_T\left[\frac{W_T^{1-\gamma}}{1-\gamma}\middle| no\ policy\ change\right], E_T\left[\frac{CW_T^{1-\gamma}}{1-\gamma}\middle| policy\ change\right]\right\} \quad (4)$$

where $W_T = B_T = \int_0^1 B_T^i d_i$ means the value of aggregate capital and C denotes the political cost the government incurs in case a new policy is implemented. This cost is randomly drawn at any given time τ from a lognormal distribution centered at $C = 1$; this means, $c \equiv \log(C) \sim N\left(-\frac{1}{2}\sigma_c^2, \sigma_c^2\right)$.

Pástor and Veronesi (2013) claim that when investors learn the value of C at τ, the government uses it to make the policy decision. The authors refer to σ_c^2 as the 'political uncertainty'. This creates an entire element of surprise in the changes in policy the government makes resulting in stock price responses at time τ. This effect could also be expected in this investigation. Economic uncertainty embodies the non-anticipated policy changes that both stock and commodity investors are subjected to. Additionally, Pástor and Veronesi (2013) ascertain that their model captures all aspects of politics, namely redistribution, special interests, corruption, etc. through the political cost C. Specifically, the researchers explain that the randomness of C reveals the challenge in predicting the outcome of the political process, which is intricate and sometimes unclear. In essence, Pástor and Veronesi (2013) offer the foundational concepts necessary to understand the influence of uncertainty on financial markets.

Now we turn to the portfolio selection model of Markowitz (1952) and to the discussions brought up in Girardi (2015) and Tang and Xiong (2012) to explain whether commodities futures might be associated with stock prices. The model of Markowitz (1952) is a mathematical framework that allows infinite combinations of assets from a portfolio and the percentages that ought to be invested in each of them to obtain the best possible outcome. Formally, the mean-variance portfolio theory follows

$$E(R_p) = \sum_i^n w_i\, E(R_i) \quad (5)$$

$$\sigma_p^2 = \sum_i^n w_i^2\, \sigma_i^2 + \sum_i^n \sum_{j\neq i}^m w_i w_j Cor_{ij} \quad (6)$$

where $E(R_p)$ denotes the expected return of the portfolio, $E(R_i)$ is the expected return on asset *i*, and w_i symbolizes the weighting of component asset *i*. Within the context of this paper, this means the proportion of stocks and commodity futures in the portfolio. In equation (6), σ_p^2 indicates the portfolio return variance, σ_i is the standard deviation of the

return on asset *i*, and Cor_{ij} represents the correlation coefficient between the returns on assets *i* and *j*. The model considers that $\sum_i^n w_i = 1$ and $w_i \geq 0$.

Equation (5) shows the expected return of the portfolio, diversified into live cattle and stocks and/or soybeans futures and stocks, for example, while equation (6) presents the portfolio risk that takes into account the correlations between these assets. In addition, to illustrate how stocks and commodities are intertwined, we briefly provide some theoretical insights into their associations.

The financialization of agricultural derivatives, for example, suggests an alternative to investors that are active in stocks and commodity markets (Girardi, 2015). Tang and Xiong (2012) state that when stocks increase in value, diversification incentives might encourage investors to move money into commodities. However, Girardi (2015) argues that what drives commodity-stock associations may be a combination of financialization and economic turmoil. During periods of uncertainty, with stock prices decreasing, investors might sell commodity derivatives causing a reduction in prices. In sum, Girardi (2015) verifies that financialized commodity markets might not produce correlations between the stock and commodity markets in 'normal' periods, but they might do so during economic stress.

All in all, this section postulates the conceptual rationale behind the association between economic policy uncertainty and the commodity-stock correlations. First, the reduced-form fashion to model policy uncertainty in Pástor and Veronesi (2013) focuses on the asset price implications of policy changes that could be mirrored in the commodity-stock markets relationship. After that, the portfolio theory of Markowitz (1952) and the brief discussions in the relevant related literature elucidates the drivers of such interactions.

3.3. Data and preliminary analysis

This paper uses daily closing prices[28] of the B3 live cattle and soybeans nearby contracts[29], and the Bovespa index between January 3, 2014 and July 10, 2023, retrieved from Barchart and Yahoo Finance, respectively. The economic policy uncertainty (EPU)

[28] Due to the impracticality of deflating daily price data, particularly given the frequency and volume of observations, as in Badshah et al. (2019), and Zhao and Wang (2022), we have opted to use nominal prices in our analysis.
[29] Nearby futures contracts refer to the futures contracts that are closest to expiration. These contracts typically have the highest trading volume and liquidity because they represent the most current delivery period.

in Brazil, the US, China, and the world are monthly data from January 2014 to July 2023. These are newspaper-based indices computed each month, counting the number of articles containing uncertainty-related terms such as uncertainty or uncertain, economic or economy, congress, legislation, regulation, deficit etc. The EPU indices for Brazil, the US, and the world are constructed by Baker et al. (2016) based on the top newspapers in these locations. Davis et al. (2019) use the same method to construct the index for China based on major mainland newspapers. These four indicators and further information are available both in table 1 and at www.policyuncertainty.com.

We are mindful that portfolio adjustments on a daily basis are often impractical due to transaction constraints and associated expenses (Zhao & Wang, 2022). Then, we follow Badshah et al. (2019), Pástor and Veronesi (2013), and Zhao and Wang (2022) by computing monthly data. Section 3.4.1 elaborates on the methodology employed in this regard.

Figure 1 plots the natural logarithm of the live cattle (BRL), soybeans (USD), and the Ibovespa (BRL) series. Live cattle and stocks mostly show an upward trend from 2014 to 2023, while the price of soybeans futures exhibits a downward trend until late 2020. Since soybeans futures are traded in US dollars, their performance could reflect not only the price of the commodity itself but also changes in the exchange rate between the Brazilian real and the American currency. All three price series evidence a positive slope in the years following the outbreak of Covid19 (between 2020 and 2021)[30]. Throughout the live cattle, soybeans, and Ibovespa series we note sudden spikes or drops which may indicate periods of increased volatility in those markets.

Figure 2 shows the raw level of EPU in Brazil, the US, China and the world. The series for Brazil shows relatively higher levels of uncertainty with respect to the world from 2014 to 2018, probably due to political stress and a presidential impeachment in 2016, for instance. EPU in the US captures evidence of specific uncertain periods such as the election of Trump from late 2016 to early 2017, and the Covid19 pandemic in 2020. EPU in China, panel (c), seems to show the trade dispute between the US and China in 2018

[30] The World Health Organization (WHO) declared the outbreak of the coronavirus as a public health emergency of international concern on Jan 30, 2020. Brazil confirmed its first case of coronavirus contamination on Feb 26, 2020. On Mar 11, 2020, the WHO declared Covid19 a pandemic (Pan American Health Organization - PAHO, 2024).

and certainly the Covid19 pandemic as from 2020. Figure 2 illustrates that uncertainty in the US is very much aligned with global uncertainty.

As an additional exercise (see appendix) we estimate the effect of uncertainty on commodity-stock dynamic correlations by replacing our EPU regressors with the Chicago Board of Options Exchange (CBOE) implicit volatility index, the VIX. This measure has been extensively referenced in the literature as an indicator of market sentiment, financial stress, and overall market uncertainty (Cesa-Bianchi et al., 2020; Silvennoinen & Thorp, 2013; Umar et al., 2021). Rising VIX values suggest that investors expect greater volatility in the near future, indicating increased fear or uncertainty.

In this study we also look at the joint effect of certain economic conditions during periods of uncertainty and how they affect the time-varying correlations between the live cattle-, soybeans-, and stock markets in Brazil. For that we use monthly information from the Instituto Brasileiro de Geografia e Estatística (IBGE) on the industrial production growth (IPG) of Brazil. These data refer to percentage changes in relation to the same period of the previous year. We also consider monthly information on the economic activity index (EAI), the Brazilian interest rate (Selic) as economic state drivers from the Central Bank of Brazil. Lastly, we incorporate the USD/BRL real effective exchange rate index (ER) from the Federal Reserve Bank of St. Louis. Table 1 summarizes all the data used in this paper.

Table 1 – List of variables and abbreviations

Variable	Abbreviation	Description	Frequency	Source
Live cattle market	$Live\ cattle$	Return of live cattle futures contracts, nearest by volume, in the Brazilian stock market	Daily	Barchart
Soybeans market	$Soybeans$	Return of soybeans futures contracts, nearest by volume, in the Brazilian stock market	Daily	Barchart
Brazilian stock market	$Ibov$	Return of Ibovespa	Daily	Yahoo Finance
Economic policy uncertainty in Brazil	EPU_{BR}	Natural log of the news-based EPU for Brazil based on the largest newspaper in Brazil: the Folha de São Paulo	Monthly	policyuncertainty.com
Economic policy uncertainty in the US	EPU_{US}	Natural log of the news-based EPU for the US based on the 10 largest newspapers in the US	Monthly	policyuncertainty.com
Economic policy uncertainty in China	EPU_{CN}	Natural log of the news-based EPU for China based on two mainland Chinese newspapers: the Renmin Daily and the Guangming Daily	Monthly	policyuncertainty.com

Economic policy uncertainty in the world	EPU_{WR}	Natural log of the news-based EPU for the world based on the GDP,ppp-weighted average of national EPU indices for 21 countries	Monthly	policyuncertainty.com
CBOE Volatility index	VIX	Constant, 30-day expected volatility of the U.S. stock market, derived from real-time, mid-quote prices of S&P500 index call and put options.	Monthly	Yahoo Finance
Industrial production growth of Brazil	IPG	% changes in relation to the same period of the previous year	Monthly	IBGE
Economic activity index of Brazil	EAI	Natural log of the index for the evolution of agriculture, industry, and service-sector products aggregated with weights derived from the Brazilian National Accounts	Monthly	dadosabertos.bcb.gov.br
Brazilian interest rate	$Selic$	% Brazilian national interest rate accumulated in the month	Monthly	dadosabertos.bcb.gov.br
Real effective exchange rate (USD/BRL)	ER	Real broad effective exchange rate for Brazil (2020=100)	Monthly	Federal Reserve Bank of St. Louis

Source: Authors' elaboration.

Additionally, to gain a deeper understanding of the differences between the live cattle and soybeans futures contracts traded on B3, Table 2 presents information about their structure and key characteristics.

Table 2 – Live cattle and soybeans contracts specifications

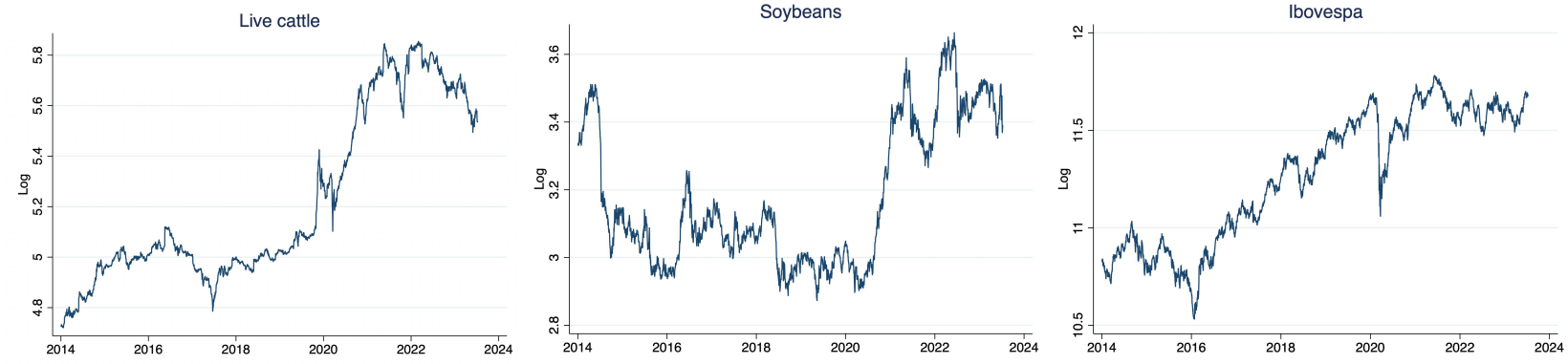
	Live cattle	Soybeans
Underlying	Male steers with at least 16 net arrobas carcass weight and a maximum age of 42 months	Cash-Settled Soybean Futures Contract at the Price of the CME Group Mini-Sized Soybean Futures Contract
Ticker	BGI	SJC
Contract size	330 net arrobas.	450 bags weighing 60-net kilograms each (corresponding to 27 metric tons).
Quotation	Brazilian Reals per net arroba to two decimal places	Dollars of United States per bag to two decimal places
Tick size	BRL 0.05	USD0.01
Round-lot	1 contract	1 contract
Last trading day	Last trading day of the expiration date	2nd business day preceding the contract month
Expiration date	Last trading day of the contract month	2nd business day preceding the contract month
Contract months	All months	January, March, May, July, August, September, and November
Settlement on expiration	Cash settlement	Cash settlement
Daily mean volume	1281	209
Daily mean price	BRL 197.38	USD 24.43

Source: B3 and authors' elaboration.

Note: Mean volume and mean prices calculated between January 3, 2024 and July 10, 2023. Numbers rounded to two decimal places.

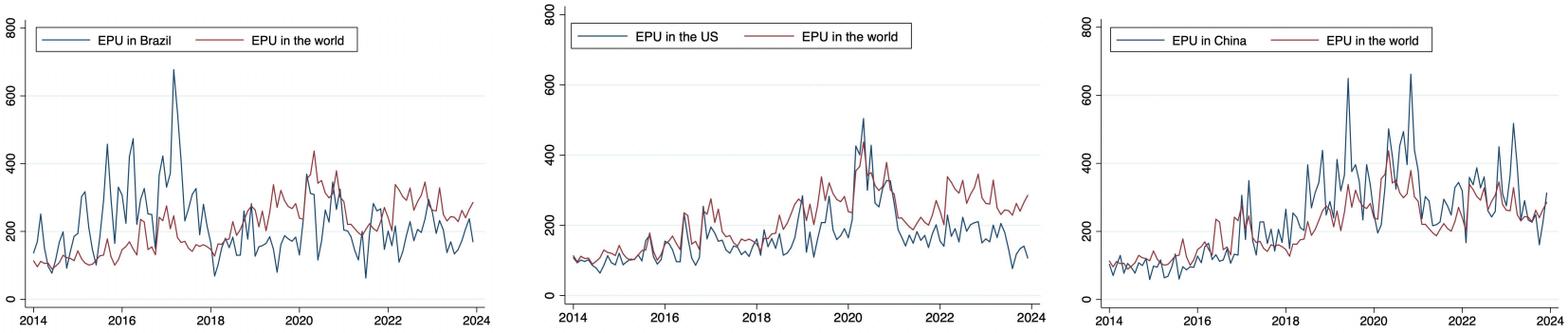
Table 2 briefly compares the broad specifications of live cattle and soybean futures contracts traded on B3, highlighting several key distinctions in contract structure, quotation, and liquidity. For this early analysis, we focus on the aspects of market liquidity, where significant differences exist between the two contracts. For instance, the daily mean volume for live cattle is reported at 1,281 contracts, with a mean price of BRL 197.38 (or approximately USD 39[31]), indicating higher trading activity compared to the soybean contract, which has a lower daily mean volume of 209 contracts and a mean price of USD 24.43. This disparity in volume may reflect different levels of market participation and investor interest, potentially stemming from varying levels of market exposure, economic significance, or commodity utility between the two futures. The higher liquidity in the live cattle market suggests more robust trading activity, which can facilitate better price discovery and lower transaction costs for traders and investors. This enhances the attractiveness of the live cattle futures contract relative to the soybean contract in terms of market dynamics and investment feasibility.

[31] For the purpose of broad comparison, the mean price of live cattle futures was roughly converted into US dollars using an exchange rate of BRL5.



Panel (a) Panel (b) Panel (c)

Figure 1 – Natural log of live cattle, soybeans futures contracts, and the Bovespa index from Jan 3, 2014 to July 10, 2023
Source: Authors' elaboration.



Panel (a) Panel (b) Panel (c)

Figure 2 – Series of EPU in Brazil, the US, China, and the world from January 2014 to July 2023
Source: Authors' elaboration.

3.4. Research design

This section elaborates on the methodology employed in this study aiming to assess whether the interactions between the live cattle-, soybeans-, and the Brazilian stock market respond to economic uncertainty from January 2014 to July 2023. First, we estimate the dynamic conditional correlations between the aforementioned commodity futures and the stock market. Then, we evaluate if economic policy uncertainty in Brazil and abroad affects these time-varying interactions.

3.4.1. DCC-GARCH model

The widespread adoption of the DCC-GARCH model stems from its effectiveness in capturing the dynamic changes in asset correlations over time. The motivation for using such a model lies in its ability to capture time-varying relationships between multiple financial assets, making it ideal for analyzing dynamic relationships, such as those between commodity and stock markets (Engle, 2002). Its advantages include flexibility in modeling changing correlations over time and accounting for volatility clustering, which is common in financial data. However, its limitations involve the complexity of estimation, computational intensity, and potential sensitivity to model specification, which may affect the results. In our investigation, we follow an array of studies that also employ a DCC-GARCH approach to explore the interactions between different asset classes, namely, Jones and Olson (2013), Kang and Yoon (2020), Mensi et al. (2013), Swamy and Padma (2020), and Wen and Wang (2021).

First, we adopt the $ARMA(p,q) - GARCH(m,n)$ model to characterize the marginal distribution of the univariate asset returns (live cattle, soybeans, Ibovespa). The lag order of the autoregressive, moving average, arch, and garch components is defined based on the AIC and BIC criteria. The return generating process is then specified as:

$$r_{i,t} = \mu_i + \theta_1 r_{i,t-1} + \cdots \theta_p r_{i,t-p} + \phi_1 \varepsilon_{i,t-1} + \cdots \phi_q \varepsilon_{i,t-q} + \varepsilon_{i,t} \quad (7)$$

$$\varepsilon_{i,t} = h_{ii,t}^{1/2} e_{i,t} \quad (8)$$

where $r_{i,t}$ denotes the daily return of asset i (live cattle, soybeans, Ibovespa) in t, μ_i indicates the unconditional mean, the parameters $\theta_1,\ldots,\theta_p$ specify the coefficients of the autoregressive terms while the parameters $\phi_1,\ldots,\phi_q$ represent the coefficients of the

moving average regressors. Finally, $\varepsilon_{i,t}$ is the innovation, $e_{i,t}$ indicates the standardized residuals generated from a normal distribution with ν_i degrees of freedom, and $h_{ii,t}$ determines the dynamic conditional variances of $\varepsilon_{i,t}$ which we obtain from the $GARCH(m,n)$ process that follows:

$$h_{ii,t} = \omega_i + a_1 e_{i,t-1}^2 + \cdots + a_m e_{i,t-m}^2 + b_1 h_{ii,t-1} + \cdots + b_n h_{ii,t-n} \quad (9)$$

where $\omega_i > 0$, $a_1,\ldots,a_m > 0$, $b_1,\ldots,b_n > 0$, and $a_i + b_i < 1$ to ascertain the stationarity of the $GARCH(m,n)$ model.

Then we follow Badshah et al. (2019) and Zhao and Wang (2022) and use the standardized residuals (standardized return shocks) to estimate the DCC-GARCH model proposed by Engle (2002). By doing that we obtain both the time-evolving variance-covariance matrix H_t and the dynamic conditional correlation matrix R_t, assumed to follow a DCC (1,1) process, as specified:

$$H_t = D_t R_t D_t \quad (10)$$

$$Q_t = (1 - \alpha - \beta)\bar{Q}_t + \alpha e_{t-1} e_{t-1}^T + \beta Q_{t-1} \quad (11)$$

$$e_{t-1} = (e_1, e_2) = (t^{-1}(u_{1,t-1}), t^{-1}(u_{2,t-1})) \quad (12)$$

$$R_t = diag(Q_t)^{-\frac{1}{2}} Q_t diag(Q_t)^{-\frac{1}{2}} \quad (13)$$

in which $D_t = diag(h_{ii,t}^{\frac{1}{2}})$ denotes the diagonal matrix containing the square roots of time-varying conditional variance from the preceding GARCH models. Q_t represents the time-varying covariance matrix of standardized residuals, while $\bar{Q}_t$ stands for the unconditional variance-covariance matrix. The coefficient α signifies the impact of the lagged standardized residual product on the dynamic conditional correlation, whereas β denotes its persistence. A stationary process commands that $\alpha > 0$, $\beta > 0$, and states that the closer $\alpha + \beta$ is to unity, the stronger the persistency of the correlation.

Again, we restate that in real-world scenarios making portfolio adjustments on a daily basis is not feasible nor significant due to transaction constraints and expenses (Zhao & Wang, 2022). Hence, our subsequent empirical analyses shift the focus to a monthly frequency. Following the approach outlined in Badshah et al. (2019), Pástor and Veronesi (2013), and Zhao and Wang (2022), we compute the monthly correlations based on the daily dynamic conditional correlations obtained from the DCC-GARCH, averaging all daily data correlations within each month.

3.4.2. Main regression

To investigate the effect of EPU on the time-varying interdependence between live cattle-, soybeans-, and stocks, we regress the averages within each month of the estimated daily dynamic conditional correlations on our measure of uncertainty. In other words, by using the estimated dynamic conditional correlations between the commodity (live cattle or soybeans) and stock markets as the dependent variable, we are able to assess whether the interactions between these two markets respond to shocks in economic policy uncertainty.

First, we evaluate the single effect of EPU in Brazil on the correlations between live cattle futures and stocks, as well as soybeans futures contracts and the Ibovespa, in line with comparable studies such as Badshah et al. (2019), Pástor and Veronesi (2013), and Zhao and Wang (2022). We repeat the same econometric exercise considering the EPU in the US, China, and the world to check if uncertainty originated abroad also affects the interactions between the commodity and stock markets in Brazil. Therefore, equation (14) exhibits our baseline regression:

$$Cor_t = \delta_0 + \delta_1 Cor_{t-1} + \delta_2 EPU_{i,t} + \xi_t \quad (14)$$

in which Cor_t represents the natural logarithm of the monthly time-varying estimated correlations between a commodity futures contract (live cattle or soybeans) and the Brazilian stock market (Ibovespa) from January 2014 to July 2023. $EPU_{i,t}$ is the natural logarithm of the economic policy uncertainty index originating in country $i = \{BR, US, CN, or\ WR\}$. As in Badshah et al. (2019) and Zhao and Wang (2022), we include in equation (14) the lagged regressor of our outcome variable in the right-hand-side to control for autocorrelation[32]. We also run equation (14) replacing the EPU regressors with an alternative measure of uncertainty, the VIX[33], which provides a snapshot of market sentiment and expectations, helping investors gauge the level of uncertainty and potential risks in the market.

Second, we look to explain the simultaneous effects of EPU in Brazil, the US,

[32] As in Badshah et al. (2019) all estimations were also performed without the lagged regressor of the dependent variable. The estimated parameters exhibited no significant variation, with diferences observed in the fourth and fifth decimal places, helping to rule out potential endogeneity concerns.

[33] Equation (14A) $Cor_t = \delta_0 + \delta_1 Cor_{t-1} + \delta_2 VIX_t + \xi_t$ exhibits the alternative regression. The VIX index is not available for the months of June 2023 and July 2023. Therefore, data on this uncertainty measure span from January 2014 to May 2023.

China, and the world on the interactions between the Brazilian commodity-stock markets. This has not been done in Badshah et al. (2019) but was estimated in Zhao and Wang (2022) for the simultaneous association between the EPU in the US and China and the oil-, gold-, and the US and Chinese stock markets. We advance on Zhao and Wang (2022) by considering not only an emerging economy as our main interest, Brazil, but also the economic uncertainty originating from major economies and trade partners of Brazil and an aggregate measure of the world's economic policy. By doing this we unveil the differences between domestic and international economic uncertainty in the commodity-stock interdependence in Brazil and see who plays the more prominent role across the market co-movements. For that we estimate the following equation:

$$Cor_t = \delta_0 + \delta_1 Cor_{t-1} + \delta_2 EPU_{BR,t} + \delta_3 EPU_{US,t} + \delta_4 EPU_{CN,t} + \delta_5 EPU_{WR,t} + \xi_t \quad (15)$$

where the regressors in equation (15) are similar to those laid out in equation (14).

Third, we attempt to understand whether the uncertainty effect is intensified under specific economic conditions. For this, we use four proxies[34], namely, industrial production growth (IPG), economic activity index (EAI), Selic, and the USD/BRL exchange rate (ER). Then we assess the joint effect of EPU for each country $i = \{BR, US, CN, or\ WR\}$ and those economic states on the pairwise interconnectedness among the Brazilian commodity-stock markets by estimating the following model:

$$Cor_t = \delta_0 + \delta_1 Cor_{t-1} + \delta_2 EPU_{i,t} + \delta_3 EC_{j,t} + \delta_4 EPU_{i,t} EC_{j,t} + \xi_t \quad (16)$$

where $EC_{j,t}$ indicates the economic condition $j = \{IPG, EAI, Selic, or\ ER\}$ in month *t* and δ_4 is our main interest in equation (16) for it captures the extent to which the relationship between economic policy uncertainty and the estimated dynamic correlations varies depending on the prevailing economic conditions. In simpler terms, it indicates whether the impact of economic policy uncertainty on the correlations between the commodity and the stock markets strengthens or weakens under certain economic states. Motivated by Pástor and Veronesi (2013), we deem that the effect of uncertainty on correlations should be more evident during periods of unfavorable economic environments, as political shocks tend to have a greater influence on asset prices when the economy is weaker. As in

[34] Badshah et al. (2019) perform a similar analysis for the US considering the AAA/BBB corporate bond default spread, the US industrial production growth, Shiller's price to earnings ratio for the aggregate US stock market, the Chicago Fed National Activity Index, and the NBER recession dummy. A reviewer kindly suggested we included the USD/BRL exchange rate in our estimations. This could be particularly interesting in the context of this research as the futures contracts of soybeans traded on B3 are quoted in USD.

equation (14), we also run equation (16) with the VIX[35].

All the regressions laid out in this section rely on time-series analyses. Using the ordinary least squares (OLS) estimator in this context could be problematic due to potential violations of the classical assumptions underlying OLS estimation. One of these assumptions is that the residuals are independently and identically distributed. However, in time-series data, observations are typically correlated over time, violating the independence assumption. For this reason, we follow Badshah et al. (2019) and apply Newey-West standard errors to address this issue by correcting for heteroscedasticity and autocorrelation, and then generate consistent estimates of standard errors in OLS regression.

Estimating OLS with Newey-West standard errors offers significant advantages. First, it addresses autocorrelation by incorporating information on the serial correlation structure of residuals. Second, it accommodates heteroscedasticity by adjusting standard errors to reflect evolving variance structures in the data. Additionally, it yields more efficient estimates of standard errors compared to traditional approaches assuming independence and constant variance. Ultimately, it ensures the validity of statistical inference in OLS regression amidst autocorrelation and heteroscedasticity. Consequently, hypothesis tests, confidence intervals, and p-values derived from OLS regression with Newey-West standard errors are more dependable in time-series analysis.

3.5. Empirical findings

This section presents the empirical findings of our analysis. First, summary statistics provide an overview of central tendencies and dispersion within the dataset, along with the estimation procedures we adopted before approximating the DCC-GARCH model. Second, we present the results of the time-varying correlations. Then, we disclose the influence of uncertainty on commodity-stock market correlations. Lastly, we report on the joint effect of policy uncertainty and economic conditions on our outcome variable.

3.5.1. Summary statistics

Table 3 presents descriptive statistics of the return series of live cattle, soybeans

[35] For our additional exercise, equation (16A) follows $Cor_t = \delta_0 + \delta_1 Cor_{t-1} + \delta_2 VIX_t + \delta_3 EC_{j,t} + \delta_4 VIX_t EC_{j,t} + \xi_t$.

and Ibovespa prices along with the summary statistics of our set of uncertainty measures and economic condition variables. In addition, figure 1A in the appendix shows the trajectories of the return series of the assets considered in this study over the sample period.

Table 3 – Summary statistics

Variable	Mean	Std. Deviation	Min	Max	Skewness	Kurtosis	JB test	Q(15)
$Live\ cattle$	0.034	0.907	-6.584	7.713	0.631	17.695	21000***	60.753***
$Soybeans$	0.002	1.347	-14.944	6.519	-1.017	13.398	11000***	42.460***
$Ibov$	0.036	1.611	-15.994	13.023	-0.919	16.082	17000***	84.734***
EPU_{BR}	5.306	0.426	4.137	6.518	-0.036	3.301	0.477^{NS}	10.293^{NS}
EPU_{US}	5.036	0.390	4.157	6.223	0.524	3.352	6.107**	14.340^{NS}
EPU_{CN}	5.350	0.568	4.076	6.495	-0.352	2.213	5.573*	21.840^{NS}
EPU_{WR}	5.304	0.383	4.493	6.080	-0.311	2.039	6.559**	19.282^{NS}
VIX	18.619	7.218	9.510	53.540	1.717	7.272	141.500***	191.85***
IPG	-1.467	6.949	-27.670	34.840	0.935	10.933	332.200***	12.118^{NS}
EAI	4.936	0.045	4.774	5.064	-0.426	4.471	14.450***	26.823**
$Selic$	0.740	0.320	0.130	1.220	-0.342	1.822	9.287***	8.942^{NS}
ER	122.449	18.323	90.520	155.160	-0.040	1.882	6.025**	786.96***

Notes: Each series of returns is calculated as $r_i = [100 * (logP_{i,t} - logP_{i,t-1})]$, where $P_{i,t}$ and $P_{i,t-1}$ indicate the price of a futures contract of asset i – live cattle (BRL), soybeans (USD), or Ibovespa (BRL) – in t and in $t-1$, correspondingly. EPU indicates the natural log of the uncertainty index, and VIX, the implicit volatility index based on S&P500. IPG in %, EAI in log, $Selic$ in %, and ER as an index (2020 = 100). The null hypothesis of the Jarque-Bera (JB) test is that the data follow a normal distribution. Q(15) is the estimated chi^2 statistic for the Ljung-Box test for autocorrelation computed at the 15^{th} lag whose null hypothesis is that the data series being tested has no autocorrelation. ***, **, and * represent statistical significance at 1%, 5%, and 10% level, correspondingly. NS indicates statistically non-significant. Numbers rounded to three decimal places.
Source: Research results.

From January 2014 to July 2023 the return series of both commodities, live cattle and soybeans, and also the return series of Ibovespa exhibit positive average returns. The standard deviation of the returns of Ibovespa (BRL 1.61), soybeans (USD 1.35), and live cattle (BRL 0.91) suggests the levels of variability around the mean. We verify that the mean values of the EPU variables are on average close to 5 with China showing a slightly higher standard deviation relative to Brazil, the US, and the world.

In table 3, the skewness values indicate the asymmetry of the return distributions. Live cattle reveal a positive skewness of 0.631, indicating a right-skewed distribution. This suggests that while most of the returns might cluster around the mean, there is a tendency for occasional larger positive returns, leading to a longer right tail in the distribution.

Conversely, soybeans and the Ibovespa exhibit negative skewness of -1.017 and -0.919, respectively, suggesting a left-skewed distribution. This entails that while the majority of returns might be clustered to the right of the mean, there are occasional larger negative returns, resulting in a longer left tail in the distribution.

The kurtosis values for live cattle, soybeans, and the Ibovespa are all positive and greater than 3, ranging from 13.398 for soybeans to 17.695 for live cattle. This excess kurtosis suggests heavier-tailed distributions with a higher likelihood of extreme observations compared to a normal distribution. The non-normality of these variables is confirmed by the significant results of the Jarque-Bera test, all at the 1% significance level, leading to the rejection of the null hypothesis of normal distribution. This finding is consistent with existing literature (Adekoya & Oliyide, 2021; Kumar et al., 2021). Furthermore, the Ljung-Box test at the 1% level indicates the presence of serial autocorrelation in all commodity and stock return series up to lag 15.

The mean value of Selic indicates that on average the Brazilian national interest rate accumulated in the month is 0.74% in the sample period. Examining the range of the data, we observe that the minimum value of IPG in Brazil is -27.7% and the maximum value is 34.8%, indicating a considerable span for this variable within the dataset. Similar to our asset variables, the Jarque-Bera test for all four economic condition regressors, namely, IPG, EAI, Selic, and ER also indicate significant departure from normality at the 1% level. The mean value for the USD/BRL real effective exchange rate index is 122.449 ranging from 90.520 to 155.160 in the January 2014 – July 2023 timespan. Except for EPU in Brazil, all our remaining economic uncertainty measures suggest that these series do not follow a normal distribution.

We conducted unit root tests, utilizing both the Augmented Dickey Fuller (ADF) and Phillips-Perron (PP) tests, on the return series of live cattle, soybeans, and the Ibovespa. Initially, these tests were applied to the natural logarithm of the futures contract price series, which confirmed that all three series exhibit non-stationarity. Subsequently, we examined the presence of a unit root in the first difference of the natural logarithm, and found that they all demonstrated stationarity at the 1% level of statistical significance. Detailed results of these unit root tests can be found in table 1A in the appendix.

In table 2A in the appendix we report the measures of unconditional correlations

among the returns of the asset variables as well as among the uncertainty and economic condition regressors. We note that the unconditional correlation between the return series of live cattle and the Ibovespa is 0.038, while the same measure for the return series of soybeans and the Brazilian stock market index is 0.148. Although this early result indicates that live cattle and the Brazilian stock markets are relatively less correlated than soybeans futures and the Ibovespa, it is important to take into consideration the probable effect of the exchange rate on the soybean price series.

Given that soybean futures are traded on B3 in US dollars, it raises suspicion that the exchange rate might serve as a transmission channel for this correlation between soybean contracts and the Ibovespa. To investigate, we compute the unconditional daily correlation between the Ibovespa and the BRL/USD exchange rate from January 3, 2014 to July 10, 2024, which reveals a coefficient of -0.76. This suggests that the relatively higher absolute correlation between soybeans and the Ibovespa is likely influenced by the exchange rate. Thus, it appears evident that the price of soybean contracts is not exogenous to the exchange rate. Moreover, the fact that these commodity contracts are quoted in US dollars potentially introduces an additional risk component between opening and closing positions. Furthermore, we highlight that these contracts face strong competition from soybean contracts traded on the Chicago exchange, where liquidity is significantly higher than on B3.

Table 3A (see appendix) presents the unconditional correlations among the uncertainty covariates. We realize that these measures for the uncertainty variable in the US and China (0.606), the US and the world (0.827), and China and the world (0.842) are high. This finding is unsurprising, considering that both the US and China are major economies that exert significant and far-reaching effects on global economic trends. The results informed in table 3A also provide insights into the relationship between economic conditions in Brazil and the country's economic policy uncertainty index. For instance, we observe that higher levels of uncertainty in Brazil are associated with lower industrial production growth, reduced economic activity, and a higher interest rate. Table 3A also evidences the unconditional correlations between our alternative uncertainty measure, the VIX index, and the the economic conditions drivers we consider in this study. Overall, we observed similar results, which entails that by using a different uncertainty measure we

find related results.

Finally, preceding the estimation of the DCC-GARCH model, we conducted tests to ascertain the presence of ARCH effects in the return series of live cattle, soybeans, and the Ibovespa, with detailed results provided in table 4A in the appendix. The outcomes revealed that the null hypothesis, positing the absence of conditional heteroscedasticity, was rejected across all assets. This gathers clear evidence of conditional heteroscedasticity in the data implying that the variance of these three series varies over time.

3.5.2. Dynamic conditional correlations

The initial step involves fitting the marginal distribution of each commodity and stock series. Table 4 shows the estimated parameters of the $ARMA(p,q) - GARCH(m,n)$ model, whose lag order was determined by AIC and BIC criteria. Panel A exhibits results of the mean equation (autoregressive and moving average terms), while panel B displays the coefficients of the variance equation (arch and garch components). The results reported in table 4 evidence that all the estimated coefficients are statistically different from zero at least at the 10% level.

Table 4 – Estimation of ARMA (p,q) – GARCH (m,n) model

Panel A: Mean equation	$Live\ cattle$	$Soybeans$	$Ibov$
μ	0.034*	0.035^{NS}	0.067**
	(0.020)	(0.023)	(0.028)
θ_1	0.128***	1.563***	
	(0.022)	(0.198)	
θ_2		-0.691***	
		(0.185)	
θ_3		-0.072*	
		(0.043)	
θ_4		0.056**	
		(0.025)	
ϕ_1		-1.554***	
		(0.196)	
ϕ_2		0.697***	
		(0.177)	
Panel B: Variance equation	$Live\ cattle$	$Soybeans$	$Ibov$

ω	0.008***	0.081***	0.098***
	(0.000)	(0.017)	(0.021)
a_1	0.025***	0.155***	0.078***
	(0.001)	(0.013)	(0.008)
b_1	0.965***	0.062**	0.877***
	(0.001)	(0.024)	(0.016)
b_2		0.750***	
		(0.030)	
Degrees of freedom	5	11	4
AIC	5716.682	7747.808	8305.525
BIC	5745.483	7811.168	8328.565
LL	-2853.341	-3862.904	-4148.762
Obs.	2,345	2,345	2,345

Notes: Panel A reports the estimation of the mean equation and Panel B reports the estimation of the variance equation for the three asset returns. ***, **, * denote the statistical significance at 1%, 5% and 10% level, respectively. NS indicates statistically non-significant. Figures rounded to three decimal places. The standard errors of each estimated parameter are in parentheses.
Source: Research results.

In panel A, the $AR(p)$ estimated parameters of the mean equation denote the impact of past values of the live cattle and soybeans series on their current values whereas the $MA(q)$ coefficients capture the impact of past forecast errors of the soybeans series on its contemporaneous value. On the other hand, in panel B, the $ARCH(m)$ estimated coefficients capture the impact of past squared innovations of the live cattle, soybeans, and Ibovespa series on their current conditional variance, while the $GARCH(n)$ parameters yield the influence of the past conditional variances of these series on their current conditional variance. Overall, in the realm of the GARCH model, the parameters a and b in the variance equation present significance across all commodity and stock series at least at the 5% level. This result confirms the presence of volatility clustering and persistence effects with the series of live cattle, soybeans, and Ibovespa.

Table 5 shows the results of the DCC-GARCH model. As in Badshah et al. (2019) and Zhao and Wang (2022), it presents the estimated coefficients of the arch and garch terms for the live cattle, soybeans, and Ibovespa standardized residual series (obtained from the $ARMA(p,q) - GARCH(m,n)$).

Table 5 – Estimation of DCC-GARCH model

	$Live\ cattle - Ibov$	$Soybeans - Ibov$
$\alpha_{commodity}$	0.025***	0.100***
	(0.004)	(0.013)

$\beta_{commodity}$	0.965***	0.888***
	(0.005)	(0.015)
Intercept	20.644***	72.145***
	(4.306)	(22.807)
α_{Ibov}	0.078***	0.076***
	(0.011)	(0.011)
β_{Ibov}	0.877***	0.880***
	(0.018)	(0.017)
Intercept	121.843***	119.103***
	(28.721)	(27.746)
$\lambda_{commodity}$	0.012^{NS}	0.012^{NS}
	(0.014)	(0.009)
λ_{Ibov}	0.755***	0.907***
	(0.221)	(0.051)
Degrees of freedom	11	11
AIC	49111.87	50364.50
BIC	49175.23	50427.86
LL	-24544.94	-25171.25
Obs.	2,345	2,345

Post-estimation statistics

Arch effect test	chi^2 test stat	p-value	chi^2 test stat	p-value
lag(5)	0.279	0.998^{NS}	0.045	1.000^{NS}
lag(15)	7.340	0.948^{NS}	0.330	1.000^{NS}
lag(25)	8.666	0.999^{NS}	0.356	1.000^{NS}
JB test	chi^2 test stat	p-value	chi^2 test stat	p-value
$Resid^2$	2000000	0.000***	42000000	0.000***

Notes: The subscript *commodity* denotes either live cattle or soybeans. ***, **, * indicate the rejections of the null hypothesis at 1%, 5% and 10% significance levels. NS indicates statistically non-significant. JB specifies the Jarque-Bera test for normality. Coefficients rounded to three decimal places. The standard errors of each estimated parameter are in parentheses.
Source: Research results.

We observe that all commodity-stock asset pairs are statistically different from zero at 1% significance level and the sums of α and β are all greater than 0.95, showing that there exists a considerable time-varying dependence process between the commodity-stock asset pairs, and the whole persistence is very strong.

Table 5 also presents post-estimation statistics aiming to verify whether the estimated DCC-GARCH model captured the arch effect. For that, we first run the Portmanteau test, which indicates the presence of ARCH effects if autocorrelation is

detected in the residuals. The consistent nonrejection of the null hypothesis across all residual series suggests that the estimated model effectively captures the dynamic properties of our data, rendering higher-order models unnecessary. Then, we run the Jarque-Bera test to check the normality assumption of the residuals. The results strongly rejected the null hypothesis meaning that the residuals do not follow a normal distribution.

To gain a deeper insight, table 6 displays the summary statistics concerning time-varying correlations.

Table 6 – Summary statistics of daily dynamic conditional correlations

	$Live\ cattle - Ibov$	$Soybeans - Ibov$
Mean	0.037	0.153
Standard deviation	0.019	0.032
Min	-0.079	-0.043
Max	0.263	0.359

Note: Figures rounded to three decimal places.
Source: Research results.

In table 6, we observe that the average time-varying correlations between live cattle and stocks are lower than those between soybeans and the Ibovespa. This insinuates that live cattle could be regarded as a better diversification option compared to soybeans. However, it is important to once again consider the potential effect of the exchange rate on the estimated mean value of the dynamic conditional correlations, given that soybean futures traded on B3 are quoted in US dollars. These correlations likely reflect not only the relationship between soybean futures contracts and the Ibovespa but also the influence of the BRL/USD exchange rate, embedded within soybean futures, on the Brazilian stock market. The price of this commodity in the Brazilian currency is defined in a way that, upon conversion, aligns with the international market. Consequently, regardless of the currency unit used (whether BRL or USD), the price of soybeans probably remains intertwined with the exchange rate.

Moreover, assuming a long position in both soybeans and the Ibovespa, we note that the standard deviation of the dynamic conditional correlations between soybeans and the stock market is relatively higher, indicating greater risk for investors. This means that investors are exposed to the risk of a decrease in commodity prices and also to the risk of a downturn in the stock market. However, if an investor were short in soybeans, for instance, betting on a decline in its prices, the interpretation of this result would be

different. In such a scenario, the relationship between the soybeans contracts and the stock market would indicate reduced risk for the investor short in commodities when the correlation with the stock market is higher.

According to table 6, the correlation of 0.037 between live cattle futures contracts and the Ibovespa, for instance, implies a very weak positive relationship between the two assets. In practical terms, this correlation indicates that the movements in the prices of live cattle futures and the Ibovespa tend to have little to no influence on each other. For an investor with long positions in both assets, this correlation implies that the risk exposure associated with fluctuations in the price of live cattle futures is largely independent of movements in the Ibovespa. Therefore, changes in the value of one asset are unlikely to significantly impact the other. As a result, the investor may benefit from diversification, as the weak correlation between the two assets helps to spread risk across different sectors of the market.

The evolutions of the dynamic conditional correlations among commodity-stock asset pairs are depicted in figure 3, enabling the observation of their time-varying nature.

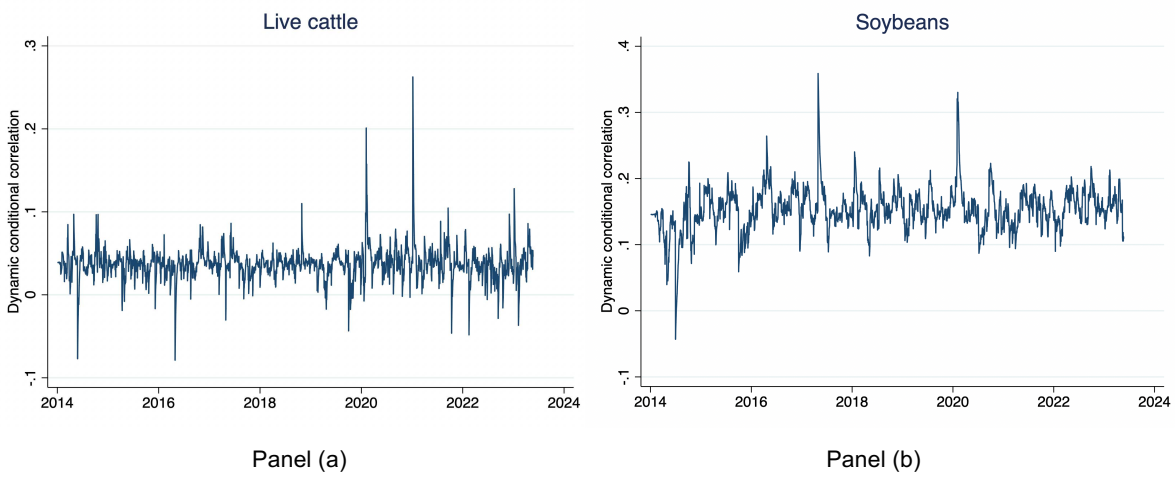


Figure 3 – Daily dynamic conditional correlations of live cattle-, soybeans- and stocks from Jan 3, 2014 to Jul 10, 2023
Source: Authors' calculations.

Notably, in figure 3 panel (a), the correlation between the live cattle market and stocks stands out. Initially mostly ranging between 0 and 0.1 until 2020, it surged to nearly 0.2 during the onset of the Covid19 pandemic. Subsequently, the time-varying interactions between live cattle and the Brazilian stock market reached its highest correlation, still

during the pandemic in 2021. In figure 3 panel (b), the correlations between soybeans and the Brazilian stock market were also impacted by the occurrence of the coronavirus, surpassing the value of 0.3. With the exception of late 2014 and late 2017, the evolving interactions between soybeans and the Ibovespa remained predominantly around its mean.

All in all, in figure 3 we verify that although the estimated dynamic conditional correlations from Jan 3, 2014 to Jul 10, 2023 are positive, they are very low, demonstrating that both live cattle and soybean futures contracts nearly do not affect (or are affected by) the Brazilian stock market given the weak nature of the associations among them[36]. This result may be attributed to different factors. The liquidity of the Brazilian stock market as a whole, coupled with the volume of contracts traded, could contribute to the independence of live cattle (and/or soybeans) futures contracts and the Ibovespa. Additionally, the influence of the exchange rate as an additional risk factor and its possible effect on the interactions between soybean contracts and the Ibovespa could further explain the relatively higher correlation we observed (in respect of live cattle).

3.5.3. The effect of uncertainty on the correlations between commodity and stock markets

Now we head to the main results of this study; that is, the influence of economic policy uncertainty on the dynamic conditional correlations between the commodity and stock markets in Brazil. Table 7 shows the results of the single (model 1) and simultaneous (model 2) effect of EPU on the correlations between the commodity and stock markets. This means that model 1 evaluates the effect of EPU either originated in Brazil, the US, China, or the world on the correlations between live cattle (or soybeans) and the Brazilian stock market. On the other hand, model 2 examines how EPU in these four regions affect our outcome variable concurrently.

Table 7 – Estimation for the single (model 1) and simultaneous (model 2) effect of policy uncertainty on commodity-stock correlations

[36] In addition to the DCC-GARCH model, we also estimated the conditional correlations using the BEKK-GARCH model, employing residuals derived from an ARMA model. The correlations estimated through the BEKK-GARCH were similarly low on average, with slight differences in magnitude compared to those obtained via the DCC-GARCH model. These nuances could stem from the structural differences between the two models. However, in line with the related literature, in this study we opted for the DCC approach.

Variables	$Live\ cattle - Ibov$				
	Model 1: Eq. (14)				Model 2: Eq. (15)
EPU_{BR}	0.125***				0.078***
	(0.034)				(0.028)
EPU_{US}		0.203***			0.448***
		(0.026)			(0.105)
EPU_{CN}			-0.008^{NS}		-0.123^{NS}
			(0.030)		(0.094)
EPU_{WR}				0.085^{NS}	-0.156^{NS}
				(0.059)	(0.246)
Intercept	-3.969***	-4.337***	-3.226***	-3.725***	-4.556***
	(0.305)	(0.177)	(0.266)	(0.402)	(0.512)
F-stat	13.57	30.42	0.22	2.26	36.74
Prob>F	0.000	0.000	0.8004	0.1093	0.000
Obs.	114	114	114	114	114

Variables	$Soybeans - Ibov$				
	Model 1: Eq. (14)				Model 2: Eq. (15)
EPU_{BR}	0.081**				0.069**
	(0.031)				(0.029)
EPU_{US}		0.098**			0.028^{NS}
		(0.034)			(0.031)
EPU_{CN}			0.038*		-0.036*
			0.020		(0.021)
EPU_{WR}				0.101**	0.121**
				(0.035)	(0.053)
Intercept	-1.663***	-1.772***	-1.393***	-0.035***	-2.332***
	(0.154)	(0.191)	(0.121)	(0.195)	(0.346)
F-stat	60.68	54.91	57.65	64.71	109.05
Prob>F	0.000	0.000	0.000	0.000	0.000
Obs.	114	114	114	114	114

Notes: Model 1 presents the estimated coefficients of equation (14): $Cor_t = \delta_0 + \delta_1 Cor_{t-1} + \delta_2 EPU_{i,t} + \xi_t$, and model 2 shows the estimated parameters of equation (15): $Cor_t = \delta_0 + \delta_1 Cor_{t-1} + \delta_2 EPU_{BR,t} + \delta_3 EPU_{US,t} + \delta_4 EPU_{CN,t} + \delta_5 EPU_{WR,t} + \xi_t$. Both models include Cor_{t-1} to control for autocorrelation and warrant the robustness of our estimates. ***, **, * represent the rejections of the null hypothesis at 1%, 5% and 10% significance levels. NS indicates statistically non-significant. The significance of the coefficients is based on Newey-West standard errors with 100 lags[37]. The standard errors of each estimated parameter

[37] We follow Badshah et al. (2019) and test the significance of the estimates using Newey-West standard errors based on various lags and obtain similar estimated coefficients. Greene (2000) suggests that the maximum lag L should be predetermined to ensure it is sufficiently large, such that autocorrelations beyond L are negligible and can be disregarded. While Newey-West standard errors are robust to autocorrelation, we are aware that using too many lags can introduce inefficiencies and potentially bias our estimates. However, using fewer lags resulted in statistically non-significant parameters with no change in the magnitude of the coefficients.

are in parentheses. Figures rounded to three decimal places.
Source: Research results.

First we evaluate the results of models 1 and 2 for the effect of EPU on the correlations between live cattle and the Ibovespa. The F-stat allows us to confirm that in model 1, economic policy uncertainty in China and in the world does not seem to impact how the live cattle and stock markets interact in Brazil. However, we verify a positive and statistically significant influence of uncertainty originated in Brazil and in the US on the time-varying dynamics between live cattle and the Brazilian stock market. This result indicates that as economic uncertainty arises in Brazil and in the US the interaction between the price of live cattle futures contracts and the Bovespa index becomes more correlated.

This result implies that during periods of economic uncertainty in Brazil and the US, the price dynamics of live cattle futures contracts become more closely linked to the performance of the Bovespa index. This increased correlation suggests that investors may view these markets as slightly more interconnected in times of uncertainty, potentially indicating a shared response to broader economic conditions. As economic uncertainty rises, investors might react similarly across asset classes, leading to synchronized price movements.

While our statistical analysis indicates a positive relationship between uncertainty and the correlations between a commodity (either live cattle or soybeans) and the stock market, a more nuanced economic perspective is required. This is primarily because the estimated effect of uncertainty on these correlations is exceedingly small, bordering on zero, across both models 1 and 2. Consequently, we argue that from an absolute standpoint, such an effect is negligible. Our findings not only reveal the presence of a very weak relationship between the assets examined in this study, as evidenced by the estimated dynamic conditional correlations (see table 6 and figure 3), but also demonstrate a near-null effect of uncertainty on the connections between live cattle-, soybeans-, and the stock market, as depicted in table 7.

Illustratively, in figure 3, spanning over a period of more than nine years, there were only four instances where correlations exceeded 0.1. Building on this, as we delve deeper into our analysis and investigate the impact of domestically and internationally originated EPU on the time-varying dynamics of these markets in Brazil, we unearth coefficients that

hover around zero. This likely shows that there is scant credible reason, based on the insights generated in this study, to prompt investors to seek alternative assets during periods of heightened uncertainty. In other words, while the average effect may be statistically different from zero, it does not appear to exert an economically significant influence, as shown in table 7.

We unveil a similar result for the simultaneous impact of economic policy uncertainty on the time-evolving conditional interactions between live cattle futures and the Ibovespa. In table 7, model 2 shows these estimates. Again, we confirm that uncertainty starting in China and the world has no apparent influence on how these two assets correlate in Brazil. Nevertheless, the results of model 2 are in line with model 1 stating that there exists a positive yet very small statistically significant effect running from economic uncertainty in Brazil and in the US on the correlations between live cattle futures and the Brazilian stock market. Moreover, we note that the coefficients in both models 1 and 2 entail that uncertainty in the US produced a comparatively more powerful effect than that one initiated in Brazil on the commodity-stock correlations.

The finding of the relatively more widespread impact of the EPU in the US on our dependent variable aligns with previous research examining global markets (Agnello et al., 2020; Qin et al., 2020) and correlations within stock markets (Li and Peng, 2017). Correspondingly, Zhang et al. (2019) delve into the influence of EPU emanating from China and the US on global markets. Despite China's rising influence, their research underscores the enduring dominance of the US across all market domains.

Now we turn to the results of models 1 and 2 for the effect of EPU on the correlations between soybeans and the Ibovespa. These results are also reported in table 7. The F-stat tells us that all our single EPU regressions are statistically different than zero and confirm a small positive relationship running from economic uncertainty in Brazil, the US, China, and the world to the associations between soybeans futures contracts and the Ibovespa. The estimated coefficients disclose that as these countries experience more episodes of economic uncertainty soybeans futures become a bit more correlated with the stock market in Brazil. Furthermore, we note that this relationship is the strongest for the effect of EPU in the world (0.101) followed by the US (0.098), Brazil (0.081), and China (0.038) on the cross-asset interactions in Brazil. Similar to live cattle, the change in

correlations is notably modest, keeping them at markedly low levels. These coefficients imply a rather inelastic association; for instance, a 1% increase in economic uncertainty in Brazil augments the correlations between soybeans and stocks in 0.08%.

When comparing the results of model 1 with those of model 2 regarding the response of soybeans-stock correlations to economic uncertainty, we observe slight changes. In model 2, we see that by regressing the dynamic correlations between soybeans futures and the Ibovespa on all EPU measures simultaneously, the estimated coefficients denote that there exists no effect running from EPU in the US on our dependent variable. Furthermore, when we allow economic frictions in all these four regions to vary concurrently, the EPU in China shows a different behavior toward how soybeans futures and the Brazilian stock market correlate. In model 2, we unveil that uncertainty in the Chinese economy makes these two assets less correlated.

Our analysis uncovers that uncertainty in the Chinese economy tends to decrease the correlation between soybeans futures and the Ibovespa. Despite unveiling a coefficient of -0.036 in table 7 to represent this relationship, its magnitude is exceptionally small, close to zero. One plausible explanation for this phenomenon could also be the influence of the exchange rate channel. Initially, uncertainty in the Chinese economy may impact the RMB/USD exchange rate. This fluctuation in turn affects the USD/BRL exchange rate and subsequently influences the Ibovespa. Therefore, while the coefficient suggests a negative impact of Chinese economic uncertainty on the correlation between those assets, its limited effect size implies that the exchange rate mechanism may play a significant role in attenuating this relationship.

Our motivation to run such a simultaneous model comes from Zhao and Wang (2022). Nonetheless, our unconditional correlation matrix in table 3A (in the appendix) indeed suggests that the EPU in the US, China, and the world reveal correlations ranging from 0.60 to more than 0.80. Hence, this regression model may attribute the effects of one correlated variable to the other, resulting in coefficients that do not reflect their true impact on the dependent variable.

All in all, focusing on the results of model 1, table 7 confirms the existence of a positive and statistically significant (except for EPU_{CN} and EPU_{WR} in our live cattle-stock analysis) influence of economic policy uncertainty on the interactions between the live

cattle-, soybeans-, and stock markets in Brazil from Jan 2014 to Jul 2023. Our results are somewhat similar to those found in Badshah et al. (2019) for the correlations between these commodities and the S&P500. These authors find that EPU in the US positively affects the conditional correlations between live cattle futures (and soybeans) and the stock market in the US.

In summary, in a context where uncertainty positively influences correlations between live cattle and stocks, as well as soybeans and stocks, heightened uncertainty could prompt investors to seek refuge in perceived safe havens such as live cattle and soybeans futures, potentially amplifying correlations. Nevertheless, we posit that this scenario would only be the case if, for instance, the estimated parameters in table 6 for the effect of EPU on commodity-stock correlations were significantly larger, say, three or four times bigger. Therefore, the results of this study indicate that even as economic frictions increase both in Brazil and/or abroad, we observe very little, practically null influence of uncertainty on the interactions between the markets we explore.

We argue over the possibility that other factors, such as market sentiment, geopolitical events, or specific sector dynamics, may exert a more significant impact on market correlations than economic uncertainty alone. Most importantly, we conjecture that the intrinsic characteristics of the agricultural futures market in Brazil provide a robust rationale for our findings. The limited number of commodity futures contracts traded on B3, particularly soybeans, and the lack of liquidity compared to assets negotiated on the Chigago exchange, for instance, indicate that investors often rely on alternative market fundamentals to mitigate the effects of uncertainty, thereby dampening its influence on the interactions of the commodity and stock markets in Brazil.

The results we report in table 7 are somewhat comparable to those observed in the literature. Overall, we observe a significant yet small positive effect of policy uncertainty on the time-varying correlations. This aligns with findings by Fang et al. (2018), who demonstrate that policy uncertainty significantly enhances the long-run correlation between oil and stock markets. In regards to a more similar approach, the results in Badshah et al. (2019) for the United States unveil a null effect of the EPU in the US on the correlations between live cattle and the S&P500, and a very small positive influence of uncertainty in America on the soybeans-S&P500 time-varying interactions. However, it

is important to note that these studies either considered somewhat different types of assets or different markets.

Finally, we highlight that as an additional exercise we also employ an alternative measure of uncertainty, the VIX, which measures the expectation of future volatility of the market based on the S&P500 index options. This index denotes a standard and trusted proxy for volatility in academic research, making it a valuable tool for understanding and analyzing market dynamics. Table 5A, in the appendix, shows that our additional exercise confirmed the consistency of our findings, revealing similar results for the effect of uncertainty on commodity-stock correlations. The results unveiled a positive yet minimal effect of the VIX onto the live cattle-stock (0.011) and the soybeans-stock (0.004) correlations. This strengthens the reliability of our initial results, demonstrating that the relationship between uncertainty and the correlations between commodities and stocks holds true when an alternative measure of market volatility is employed.

Hence, our primary finding underscores that even amidst periods of political uncertainty, while this variable does indeed positively impact our estimated dynamic correlations, its influence remains insufficient to escalate them to a degree necessitating portfolio adjustments or rendering the utilization of such derivatives for diversification impractical.

3.5.4. The joint influence of economic policy uncertainty and economic conditions

This section delves into the impact of uncertainty and the economic state on the correlation between commodities and stocks. During periods of political crisis, capital markets tend to experience increased volatility (Chau et al., 2014) while the uncertainty effect on asset correlations is often more prominent under hostile economic conditions (Pástor & Veronesi, 2013). Therefore, in table 8 we report the results of the joint effect of economic policy uncertainty and four different economic condition measures, namely, the industrial production growth in Brazil, the country's economic activity index, its interest rate, and the USD/BRL exchange rate on the commodity-stock correlations obtained from the DCC-GARCH model. Panels A and B in table 8 show the estimated parameters for the interaction terms in equation (16); in other words, the joint effect of EPU and economic conditions on live cattle-stock and soybeans-stock correlations, respectively.

The coefficients in panel A, for instance, symbolize the additional change in the interconnectedness between live cattle futures contracts and the Brazilian stock market associated with the interaction between EPU and economic conditions. In the case of IPG, it tells us how the relationship between uncertainty and commodity-stock correlations varies depending on the industrial production growth rate in Brazil. The negative small sign indicates that as the IPG increases, the effect of economic uncertainty in Brazil, China, and the world on the dynamic correlations becomes slightly more negative. This means that the influence of uncertainty on the associations between live cattle futures and the Ibovespa weakens very little little as Brazil's industrial production experiences growth.

A similar result is found when we considered the Brazilian economic activity index as a proxy for the country's economic performance. All the estimated parameters yield negative and statistically significant results at least at the 10% level meaning as the economic activity index increases, the effect of economic policy uncertainty (in Brazil, the US, China, or the world) on the dynamic correlations between live cattle futures contracts and the Ibovespa decreases. This means that when economic activity is high, the influence of uncertainty on the correlation between these assets is notably reduced. Put differently, akin to instances of industrial production growth, during periods of more responsive economic activity, uncertainty affects the correlations between live cattle futures and the stock market more negatively.

Badshah et al. (2019) reported comparable findings in their research, which explored the impact of EPU in the US on the correlations between live cattle futures and the S&P500. They identified a negative association between these variables when interacted with US industrial production growth and the US economic activity index. Our study extends this analysis to the Brazilian context and confirms a similar effect. Furthermore, our findings indicate that this impact is amplified when considering EPU at a global level (-3.153), as well as specifically in the US (-1.920), China (-1.914), and Brazil (-1.665).

Table 8 – Estimation of the joint influence of policy uncertainty and economic conditions on commodity-stock correlations

	EPU_{BR}				EPU_{US}				EPU_{CN}				EPU_{WR}			
Panel A	IPG	EAI	$Selic$	ER	IPG	EAI	$Selic$	ER	IPG	EAI	$Selic$	ER	IPG	EAI	$Selic$	ER
Live cattle																
$Ibov$	-0.013**	-1.665*	-0.115^{NS}	0.003^{NS}	-0.008^{NS}	-1.920*	-0.120^{NS}	0.003*	-0.020***	-1.914***	0.116^{NS}	-0.003**	-0.026***	-3.153***	-0.135^{NS}	0.001^{NS}
	(0.006)	(0.967)	(0.100)	(0.004)	(0.007)	(1.044)	(0.153)	(0.002)	(0.003)	(0.541)	(0.103)	(0.001)	(0.007)	(0.829)	(0.115)	(0.004)
Intercept	-3.896***	-45.711^{NS}	-4.494***	-1.660^{NS}	-4.142***	-53.154*	-4.802***	0.483^{NS}	-2.743***	-51.613***	-2.338**	-4.172***	-3.157***	-85.723***	-4.160***	-2.397^{NS}
	(0.246)	(28.469)	(0.599)	(1.834)	(0.144)	29.152	(0.483)	(1.242)	(0.210)	(15.509)	(0.811)	(0.661)	(0.278)	(23.452)	(0.592)	(1.879)
F-stat	13.71	13.51	15.58	11.84	39.53	20.35	46.97	161.93	33.86	28.37	1.19	3.10	30.71	17.22	5.63	5.46
Prob>F	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.3173	0.0186	0.000	0.000	0.000	0.000
Obs.	114	114	114	114	114	114	114	114	114	114	114	114	114	114	114	114
	EPU_{BR}				EPU_{US}				EPU_{CN}				EPU_{WR}			
Panel B	IPG	EAI	$Selic$	ER	IPG	EAI	$Selic$	ER	IPG	EAI	$Selic$	ER	IPG	EAI	$Selic$	ER
Soybeans																
$Ibov$	0.003^{NS}	2.476**	0.193**	0.003*	0.004**	2.231***	0.143**	0.005**	-0.001^{NS}	2.139***	0.075^{NS}	0.003*	0.002^{NS}	2.440***	0.077^{NS}	0.005*
	(0.003)	(0.793)	(0.082)	(0.002)	(0.002)	(0.401)	(0.059)	(0.002)	(0.002)	(0.253)	(0.052)	(0.002)	(0.003)	(0.398)	(0.055)	(0.003)
Intercept	-1.698***	65.948**	-0.920**	0.483^{NS}	-1.854***	57.786***	-1.597***	0.640^{NS}	-1.382***	59.578***	-1.169***	0.506^{NS}	-1.837***	66.512***	-1.641***	1.249^{NS}
	(0.179)	(22.811)	(0.298)	(1.242)	(0.208)	(10.792)	(0.131)	(1.101)	(0.176)	(7.410)	(0.326)	(1.045)	(0.208)	(10.913)	(0.230)	(1.769)
F-stat	103.03	130.48	125.16	161.93	60.26	133.79	30.08	111.63	38.01	195.59	61.40	50.86	48.90	170.02	81.41	105.43
Prob>F	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000
Obs.	114	114	114	114	114	114	114	114	114	114	114	114	114	114	114	114

Notes: Table 8 presents the estimated interaction coefficients of equation (16): $Cor_t = \delta_0 + \delta_1 Cor_{t-1} + \delta_2 EPU_{i,t} + \delta_3 EC_{j,t} + \delta_4 EPU_{i,t} EC_{j,t} + \xi_t$, that is, δ_4. In equation (16) $i = \{BR, US, CN, or\ WR\}$, and $j = \{IPG, EIA, Selic, or\ ER\}$. All regressions include Cor_{t-1} to control for autocorrelation and ensure robustness. ***, **, * represent statistical significance level at 1%, 5% and 10%. NS indicates statistically non-significant. The significance of the coefficients is based on Newey-West standard errors with 100 lags. The standard errors of each estimated parameter are in parentheses. Figures rounded to three decimal places. Source: Research results.

Turning now to the analysis of panel B in table 8 we identify the joint influence of uncertainty and economic states on the interactions between futures contracts of soybeans traded on B3 and the Brazilian stock market. We verify that changes in economic uncertainty originated in all countries (except for the US though the statistically significant coefficient is very close to zero) and industrial production growth do not impact the correlations between soybeans futures and the Ibovespa. In this case, the relationship between these variables remains relatively stable regardless of joint fluctuations in economic uncertainty and industrial production.

The stability observed in the relationship between the aforementioned variables highlights a significant aspect of the soybeans market. As indicated by the estimated parameters in panel B of table 8, the presence of statistically insignificant coefficients insinuates that international factors might exert a more pronounced influence on the soybeans market compared to domestic economic conditions like the EAI and the IPG. This difference becomes more outward when juxtaposed with the live cattle futures market, where the Selic rate exhibits a discernible impact when interacted with uncertainty in both Brazil and the US. This finding can be rationalized by the fact that soybeans contracts are quoted in US dollars, thereby establishing a direct link between the Selic rate and the exchange rate.

Conversely, we argue over the fact that live cattle futures contracts, being quoted in BRL, exhibit a relatively higher susceptibility to domestic economic conditions in Brazil. This observation is further substantiated by comparing the estimated coefficients in panels A and B of table 8, where statistically significant differences emerge across the two commodity futures markets. In table 8, out of the 12 estimated parameters, for instance, six display contrasting statistical significance between panels A and B, elucidating the varying influences between the live cattle and soybeans futures markets in Brazil.

The assessment of the joint effect of uncertainty and the economic activity of Brazil on the correlations between soybeans futures and stocks shows a different pattern vis-à-vis the industrial production, for example. The estimated parameters are positive and statistically different than zero exposing the relationship between uncertainty in Brazil, the US, China, and the world and the time-varying associations between soybeans futures and stocks in Brazil depending on the country's level of economic activity. The estimated coefficient considering the EPU in Brazil, for instance, indicates that soybeans futures and the Ibovespa become nearly 2.5% more correlated

when the economic activity in Brazil achieves higher levels compared to those moments with lower economic performance in the country.

Unlike the results for our live cattle-stock regressions in table 8 panel A, the combined influence of uncertainty in Brazil and in the US, and the Brazilian interest rate seems to be existent, positive, and statistically meaningful. Therefore, we uncover that the association running from soybeans futures negotiated on B3 and the Brazilian stock market becomes 0.193 (EPU in Brazil) and 0.143 (EPU in the US) percentage points more correlated when there is an increase in $Selic$, probably via the exchange rate channel.

The results we report in table 8 panel B are comparable to those found in Badshah et al. (2019) although the authors produced opposing outcomes in their examination of the US context. Their study also evaluated the effects of EPU in the US on the correlations between soybeans and the S&P500. Interestingly, they discovered a non-significant association between these variables when interacted with US industrial production growth, and a negative coefficient emerged when interacted with the Chicago Fed National Activity Index, their proxy for the US economic activity index. In contrast, our research unveils a distinct effect in the Brazilian context. The differences in the results observed between the Brazilian and US commodity-stock markets in response to uncertainty shocks could stem from various factors. For instance, the structure of the Brazilian economy, the composition of its commodity exports, the investors sentiment, the volume of trade contracts (Ibovespa versus S&P500), the particular features of each contract (currency, for instance) traded on B3, and the regulatory environment, to name a few.

The estimated parameters of the interaction term between the EPU in Brazil and in the US and the ER unveil a small, positive, and statistically significant coefficient indicating that as the USD/BRL exchange rate increases, the effect of the EPU index on the dynamic correlation between soybeans and the stock market becomes marginally more positive. In other words, the impact of economic policy uncertainty originated in Brazil and in the US on these correlations strengthens as the Brazilian Real depreciates relative to the US dollar.

The results reported in table 8 confirm the perception laid out in Pástor and Veronesi (2013) that there exists a pronounced impact of uncertainty on cross-market correlations amid challenging economic environments. Furthermore, our estimated coefficients mostly endorse the observation in Chau et al. (2014) that capital markets

often undergo intensified volatility in times of political stress, captured by the four economic drivers we considered in equation (16).

Finally, we conduct an additional exercise by replacing the EPU regressors with the VIX index. We report our findings in table 5A in the appendix. This analysis confirms the observed effect running from uncertainty to commodity-stock correlations under certain economic conditions. The estimated parameters in tables 8 and 5A disclose similar associated directions and a negligible effect of uncertainty on the outcome variable as IPG, EAI, Selic, and the ER increase.

Overall, we deem the market for live cattle futures contracts traded on B3 holds greater significance with respect to soybeans. This is likely due to the relatively higher number of contracts traded daily in the Brazilian stock market and the fact that these contracts are denominated in the local currency, thus shielding investors from exchange rate risk. In contrast, the soybeans market is comparatively smaller and probably influenced by the exchange rate as it is traded on B3 in US dollars. In addition, we highlight that this commodity exhibits lower daily liquidity on average among other agricultural derivatives traded on the Brazilian exchange and probably faces competition from its counterparts traded on the Chicago exchange.

In this context, the results of a study of this nature, which considers two agricultural commodities with different attributes related to their markets, shed light on the relationship between economic uncertainty and how the commodities market interacts with the stock market in Brazil. The consideration of futures contracts for two assets, with differing levels of liquidity on B3 and traded in different currencies, allows us to compare, to some extent, whether these characteristics affect the possible relationship between uncertainty and investment in the Brazilian market. Despite these differences, we unveil that overall, the agricultural derivatives market in Brazil trades a reduced volume of futures contracts on average compared to the US market, for example, which could support the notion that economic uncertainty does not significantly impact these interactions in Brazil. Therefore, we reckon that investors in the country most likely seek alternative means of risk mitigation and portfolio diversification.

3.6. Final remarks

Particularly after the subprime crisis in 2008 a lot of attention has been devoted to political and economic stability across markets. In spite of the growing strand of literature looking to understand how economic policy uncertainty might affect financial markets, few studies have explored how interactions between the commodity and stock markets in a developing emerging economy such as Brazil respond to uncertainty shocks originated domestically and internationally. Hence, in this paper we apply the DCC-GARCH model to first estimate the dynamic conditional associations running from both live cattle and soybeans futures to the Brazilian stock market. After that, we adopt the OLS estimator with Newey-West standard errors to account for autocorrelation in time series aiming to deliver reliable measures of our effect of interest. We conduct an additional exercise evaluating the joint power of uncertainty and economic states in Brazil and how they could affect the commodity-stock time-varying correlations in the Brazilian market.

This paper makes several key contributions to the literature. First, it extends the understanding of the impact of economic policy uncertainty on financial market correlations within the context of Brazil. Second, it conducts a comprehensive analysis of correlations between two relevant commodities traded on B3 (live cattle and soybeans) the the Ibovespa. Third, it provides insights into the diversification benefits of commodities for stock market investors, mainly during periods of recession-induced stock price volatility. Additionally, it uncovers the influence of uncertainty originating from foreign economies on cross-asset interactions in Brazil. Lastly, it examines whether the effect of uncertainty on commodity-stock correlations is amplified during specific economic conditions. Our findings are threefold.

As for the estimated dynamic conditional correlations, we uncover that all commodity-stock asset pairs exhibit statistical significance at the 1% level, with the sums of α and β exceeding 0.95 for each pair. This indicates a significant and robust time-varying dependence between the commodity-stock asset pairs, underscored by a strong overall persistence in their relationships. Furthermore, we discover that the live cattle DCCs demonstrated more frequent negative correlations with the Brazilian stock market compared to those of the soybeans market.

Turning to our main analysis, we identify a positive yet very small statistically significant influence of economic policy uncertainty, except for EPU_{CN} and EPU_{WR}, in

our live cattle-stock analysis. To a certain extent, we note that higher levels of uncertainty increase the correlations between soybeans futures and the stock market in Brazil. Additionally, we corroborate the existence of a marked joint effect of uncertainty on commodity-stock correlations amidst unfavorable economic conditions in the Brazilian economy. Furthermore, we perform an additional analysis by replacing the EPU variables with the VIX index, which captures implicit market volatility and denotes a common proxy for economic uncertainty.

All in all, we revisit the research questions posed in our introduction and the key insights gleaned from this study. First, we look to understand whether domestic economic policy uncertainty has substantial effects on the time-varying correlations between live cattle-, soybeans-, and the Brazilian stock market. Broadly, we confirm a positive yet very small effect of uncertainty on the dynamic conditional correlations between live cattle (soybeans) and the Ibovespa. However, we acknowledge the potential non-negligible impact of the exchange rate on our results for the soybeans market.

Second, we attempt to investigate whether economic uncertainty in the US, China, and globally influences the interaction of these commodity futures contracts with the stock market in Brazil. We uncover a similar pattern across regions, indicating a minimal effect of economic frictions on how live cattle and soybeans futures markets interact with the Brazilian stock market.

Third, we explore whether the effect of uncertainty exacerbates during specific economic conditions. Notably, due to the live cattle market's relatively higher liquidity and BRL denomination of contracts, we observe that domestic economic conditions moderately exacerbate the effect of uncertainty on live cattle-stock dynamics in Brazil. Conversely, such a pattern is not observed in the soybeans market.

In essence, our findings underscore the significance of EPU in Brazil and key foreign markets concerning the correlations of financial assets traded on B3. Furthermore, our research highlights the varying impacts of EPU across different economic conditions. We conclude that, for investors and portfolio managers operating within the Brazilian market, both live cattle and soybeans futures contracts generally function as diversifiers. Our study reveals that the interplay of uncertainty and economic conditions influences dynamic conditional correlations very little, with uncertainty tending to elevate commodity-stock correlations, particularly during recessionary periods.

The understanding that uncertainty appears to have minimal impact on how the live cattle and soybeans markets interact with stocks traded on B3 could provide significant benefits for governments and investors alike. Governments can utilize this information in marketing efforts to portray Brazil as an appealing investment destination with a range of opportunities for investors seeking to manage risks, demonstrating resilience to uncertainty shocks. For investors, our findings offer insights into developing diversification strategies that encompass both commodity futures and stocks, which do not seem very sensitive to economic frictions.

3.7. Appendix

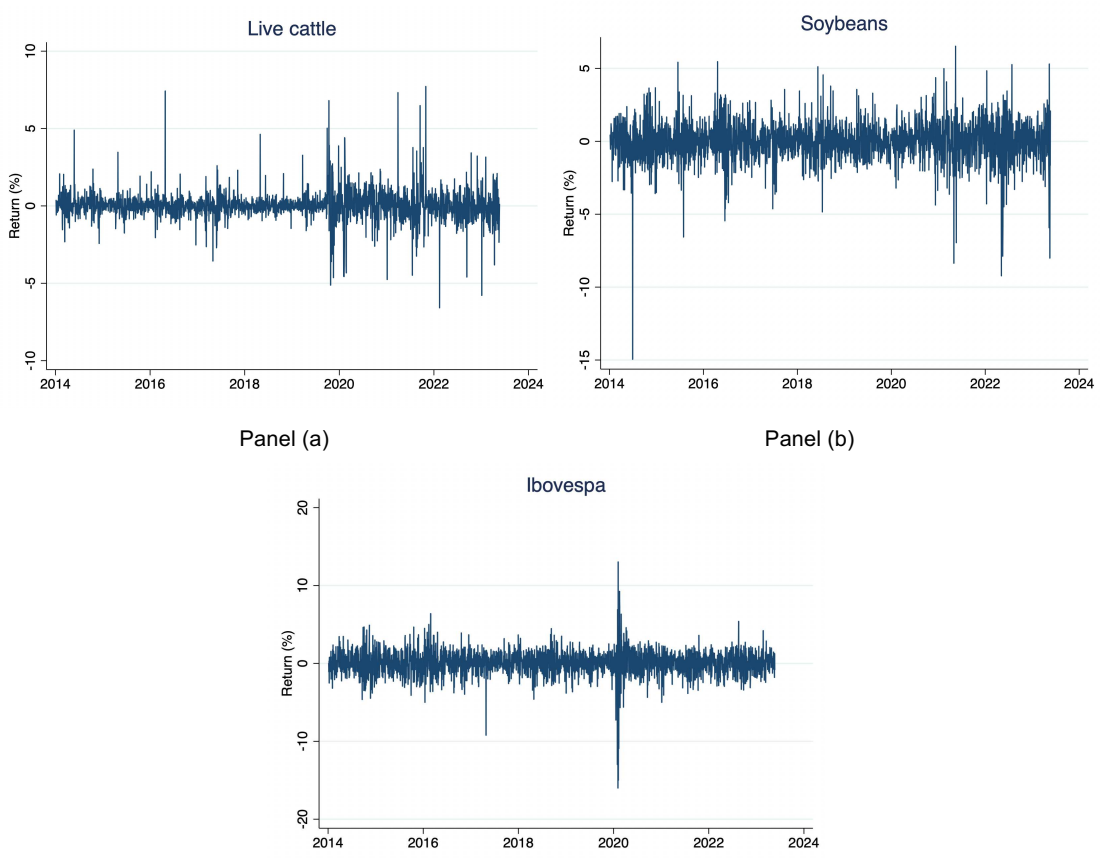


Figure 1A – Daily percentage changes of live cattle and soybeans futures contracts, and the Bovespa index from Jan 3, 2014 to July 10, 2023
Source: Authors' elaboration.

Table 1A – Unit root tests for commodity futures and the Ibovespa series from Jan 3, 2014 to Jul 10, 2023

Variables	ADF		PP	
	Test stat.	P-value	Test stat.	P-value
log($Live\ cattle$)	-1.141	0.698^{NS}	-1.184	0.681^{NS}
Δ[log($Live\ cattle$)]	-41.889	0.000***	-42.175	0.000***
log($Soybeans$)	-1.490	0.538^{NS}	-1.479	0.544^{NS}
Δ[log($Soybeans$)]	-48.114	0.000***	-48.115	0.000***
log($Ibov$)	-1.217	0.666^{NS}	-1.167	0.688^{NS}
Δ[log($Ibov$)]	-53.382	0.000***	-53.214	0.000***

Notes: ADF and PP denote the Augmented Dickey-Fuller and the Phillip-Perron tests for unit root, respectively. Δ is the first-difference operator. *** represents significance at 1% level and NS indicates statistically non-significant. Figures rounded to three decimal places.
Source: Research results.

Table 2A - Unconditional correlations among the returns series of commodity futures and Ibovespa

Variables	$Live\ cattle$	$Soybeans$	$Ibov$
$Live\ cattle$	1.000		
$Soybeans$	0.052	1.000	
$Ibov$	0.038	0.148	1.000

Source: Research results.

Table 3A - Unconditional correlations among uncertainty and economic condition series

Variables	EPU_{BR}	EPU_{US}	EPU_{CN}	EPU_{WR}	IPG	EAI	$Selic$	ER
EPU_{BR}	1.000							
EPU_{US}	0.144	1.000						
EPU_{CN}	-0.059	0.606	1.000					
EPU_{WR}	0.082	0.827	0.842	1.000				
IPG	-0.182	-0.116	0.166	0.026	1.000			
EAI	-0.249	-0.361	-0.063	-0.169	0.246	1.000		
$Selic$	0.219	-0.488	-0.508	-0.387	-0.300	0.326	1.000	
ER	0.0448	-0.5518	-0.5422	-0.6055	-0.0573	0.1168	0.414	1.000

Variables	VIX	IPG	EAI	$Selic$	ER
VIX	1.000				
IPG	-0.1266	1.000			
EAI	-0.149	0.248	1.000		
$Selic$	-0.2914	-0.3058	0.3074	1.000	
ER	-0.6141	-0.0565	0.1272	0.4243	1.000

Source: Research results.

Table 4A. Lagrange multiplier test for ARCH effects for commodity return series and Ibovespa

Variables	F-stat	P-value
$Live\ cattle$	19.290	0.000***
$Soybeans$	8.160	0.004***
$Ibov$	884.030	0.000***

Notes: *** represents significance at 1% level. Figures rounded to three decimal places.
Source: Research results.

Table 5A. Estimations using an alternative uncertainty measure

	VIX	$VIX * IPG$	$VIX * EAI$	$VIX * Selic$	$VIX * ER$
$Live\ cattle - Ibov$	0.011**	-0.001***	-0.155**	-0.045***	-0.001*
	(0.004)	(0.000)	(0.069)	(0.010)	(0.000)
Intercept	-3.575***	-3.580***	-16.098**	-4.187***	-4.720***
	(0.135)	(0.114)	(7.013)	(0.162)	(0.411)
F-stat	5.05	50.05	33.19	113.93	14.37
Prob>F	0.008	0.000	0.000	0.000	0.000
Obs.	112	112	112	112	112
	VIX	$VIX * IPG$	$VIX * EAI$	$VIX * Selic$	$VIX * ER$
$Soybeans - Ibov$	0.004**	-0.0003**	0.121^{NS}	0.000^{NS}	0.0002**
	(0.001)	(0.000)	(0.073)	(0.005)	(0.000)
Intercept	-1.778***	-1.779***	14.413^{NS}	-1.815***	-1.281**
	(0.203)	(0.209)	(9.814)	(0.265)	(0.581)
F-stat	5.57	19.060	6.40	12.170	6.25
Prob>F	0.005	0.000	0.000	0.000	0.000
Obs.	112	112	112	112	112

Notes: Table 5A shows the estimated parameters of equation (14A) $Cor_t = \delta_0 + \delta_1 Cor_{t-1} + \delta_2 VIX_t + \xi_t$ (column 1) and the coefficients of equation (16A) $Cor_t = \delta_0 + \delta_1 Cor_{t-1} + \delta_2 VIX_t + \delta_3 EC_{j,t} + \delta_4 VIX_t EC_{j,t} + \xi_t$ (columns 2 – 5). All regressions include Cor_{t-1} to control for autocorrelation. ***, **, and * represent statistical significance level at 1%, 5%, and 10%, respectively. NS denotes statistically non-significant. The significance of the parameters is based on Newey-West standard errors with 100 lags. The standard errors of each estimated coefficient are in parentheses. Numbers rounded to three decimal places.
Source: Research results.

4. CONCLUSIONS

As the world economy evolves and trade dynamics continually shift, governments worldwide may find incentives to prioritize the production of agricultural products. Additionally, following the Lehman Brothers real estate crash in the US in late 2008 and the Covid-19 pandemic in early 2020, there has been considerable focus on assessing political and economic stability across various markets.

This doctoral dissertation consists of two papers that both seek to delve into the associations between commodities and economic uncertainty. The first of them relies on China's accession to the WTO as a natural experiment to unveil the causal effect of commodity exports on institutional quality using panel data from 1997 to 2022. The second paper identifies the correlations between commodity futures and the Brazilian stock market and regresses them on the uncertainty of economic policy in Brazil and other countries using time-series information from January 2014 to July 2023.

The findings from the first study emphasize the significant impact of fluctuations in agricultural exports on a country's institutional framework. Prior research has consistently shown a negative correlation between commodity wealth and institutional development, a relationship reaffirmed in this paper concerning commodity exports and institutional quality. Moreover, through a natural experiment approach, this study establishes a causal link, demonstrating that countries prioritizing commodity-based exports witness a decline in institutional effectiveness.

In our second paper, although we statistically verify a minimal positive and significant influence of economic policy uncertainty in Brazil and in the US in the live cattle-stock analysis, we consider that, in an economic sense, this effect is rather small and should not affect investors largely. Moreover, our study underscores the diverse impact of EPU across varying economic conditions. We ascertain that, for investors and portfolio managers operating within the Brazilian market, live cattle and soybeans futures contracts typically serve as diversifiers. Our research demonstrates that the interaction between uncertainty and economic conditions, to a very small extent, influences dynamic conditional correlations.

Overall, by illuminating the nexus of commodities and uncertainty, this dissertation provides guidance for navigating the complexities of the global economy, informing strategic decisions and enhancing risk management practices.

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