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Empirical Essays on Commodity Prices

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Contents

General Introduction

Commodity prices boss the show for many developing countries. And they increased considerably over the last decade. As a result, and without surprise, commodity-producing countries growth was on average larger than in the preceding decade. World development indicators significantly improved. And the industrialization process of many areas in the world make some believe that the price increase will last. Unless this time is really different (a new black swan), commodity prices will probably decrease some day in the future, which will induce negative impacts on the economy and development of those countries. A priori, depending on commodity prices should not be different from depending on manufactured goods prices. Rising, then falling, prices of manufactured goods should make the same story for manufactured goods as the one for commodities. However, the natural resource curse literature proposes and tests some theories refuting this equivalence. Depending on commodities might be bad for growth and development in the long run.

Sachs and Warner (1995) kicked off the empirical literature by investigating on a large panel of countries the relationship between commodities and growth and found that economic dependence on oil and mineral (where the dependence is defined by the ratio of commodity exports to GDP) is correlated with slow economic growth after controlling for other attributes of the countries. Their seminal paper gave rise to a vast literature that challenges their conclusions and explores the potential channels by which the correlation occurs. We refer to Frankel (2010b) and van der Ploeg (2011) for two recent surveys.

The most important channels are the following. Firstly, commodity price inceases may give rise to the so-called Dutch disease, a phenomenon changing the structure of production through a contraction of the nonresource traded sector (typically manufacturing) and an appreciation of the real exchange rate. The Dutch disease mechanism was first modeled by Corden and Neary (1982) and Corden (1984). These models were typically developed for extractive resources but they are also applicable in other contexts (for example new high technology products) and implicitly to other commodities. Secondly, commodity prices would be highly volatile and feed more uncertainty, which affects negatively the long run growth (Aghion, Bacchetta, Rancière, and Rogoff (2009), van der Ploeg and Poelhekke (2009) and Hausmann and Rigobon (2003)). Thirdly, commodity dependence is believed to be correlated with weak institutions/governance. The idea behind this is that rents from minerals and oil are easily appropriable by governments, while for countries with no natural resource better institutions are required to coordinate the economic factors (Isham, Woolcock, Pritchett, and Busby (2003) and Collier and Hoeffler (2008)). Fourthly, natural resources lead to wars. The story is the same, natural resources such as minerals are more easily appropriable and do not require a sophisticated combination of labour and technology to be extracted (Tornell and Lane (1999) and Collier and Hoeffler (2004)). These channels, and others, help to understand those distrustful of the long run benefits of commodity-driven growth.

Considering all commodities together (and assuming they share the same features) is nevertheless certainly too simplistic in some cases. Indeed, some channels are more specific to minerals and oil than to commodities in general (institutions and wars for example). On the contrary, some channels of interaction, such as volatility and the Dutch disease, potentially affect any commodity. Since we limit our scope in this thesis on these two factors, volatility and Dutch disease, we follow a general approach and consider all commodities as a sole category, with no restriction. Our analysis in each chapter will thus include commodities, storable or unstorable, exhaustible or renewable, seasonal or unseasonal, organic or mineral. They all fall under the general definition of tradable, unprocessed or marginally processed, physical substance. Our samples will include, precious metals, agriculturals, fruits, energy.

As noted above, our scope is centered on two channels of interaction of the natural resource curse that affect a priori all commodities, without exception. We focus on the volatility of commodities in Chapters 1 and 2, especially the short run dynamic of their volatility, and the way to hedge effectively exposures on commodities by taking into account their specific volatility behaviour. In Chapters 3 and 4 we examine the Dutch disease channel. We more precisely look at the impact of commodity prices on real exchange rates of countries highly specialized in the export of one commodity. We also study the structural determinants of the real exchange rate dependence on commodity prices. Though not testing the natural resource curse itself, this thesis tends to improve the understanding of two important channels of interaction relating commodities and growth.

We show in Chapter 1 that asymmetric GARCH models, originally developed to take into account the equity leverage effect, are also relevant for modelling commodity prices. Contrary to the equity case, positive shocks turn out to be the main contributors to the conditional variance of commodity prices. Such phenomenon is an implication of the theory of storage, which relates the price of a commodity and its futures. Spot commodity prices indeed capture the tightness of the inventory, thereby feeding the volatility. We find that this inverse leverage effect, or "inventory effect", is relatively robust, for different subsamples, commodities and specifications. We also show how accounting for the inventory effect can improve risk management, through Value-at-Risk estimates.

We then illustrate in Chapter 2 the importance of considering the commodity *inventory effect* in computing optimal hedge ratios through an appropriate modelling of the variance-covariance dynamics. We show by an in-sample and out-of-sample forecast exercise that a commodity price index portfolio optimized by asymmetric BEKK-GARCH models outperforms other models. We also check robustness of our results by using a standard mean-variance utility function, which confirms the results.

Chapter 3 examines the impact that commodity prices can have on the real exchange rate of commodity exporting countries. Two recent papers show that the real exchange rate appreciates when commodity prices increase. Our analysis in this Chapter produces new estimates of this relationship by focusing on a large sample of developing countries which are specialized in the export of one leading commodity. We use non-stationary panel techniques robust to cross-sectional-dependence, and find that the price of the dominant commodity has a significant long-run impact on the real exchange rate when the exports of the leading commodity have a share of at least 20 percent in the country's total exports of merchandises. Our results also show that the larger this share, the larger the size of the impact.

Chapter 4 extends the analysis by providing new empirical evidence about the relationship that may exist between real exchange rates and commodity prices in developing countries that are specialized in the export of a main primary commodity. We more exactly investigate to which extent structural factors, such as the exchange rate regime, the degree of financial and trade openness, the degree of export concentration or the type of the commodity exports, may affect the magnitude of the effects of commodity price shocks on real exchange rates.

Our contributions to the literature on the resource curse are twofold. First, the two chapters studying the volatility of commodity prices bring tools to better understand and thus manage the risk of commodity price fluctuations. On the one hand, it shows the asymmetric behaviour of the series and illustrate the relevance of relying on asymmetric models to predict or to hedge some commodity exposures. On the other hand, our results relate the volatility dynamics to the state of the inventories. A macroeconomic policy devoted to stabilize in the short/middle run commodity prices should thus take both elements (asymmetry and importance of inventories) into account in its strategy. Second, the last two chapters, devoted to the relationship between the price of the main commodity and the real exchange rate, show the relevance of addressing the risk of Dutch disease and of fixing its implications, but also stress the importance of diversifying exports. Under a policy perspective, our results support fiscal rules and exchange rate interventions that incorporate such targets by mitigating the commodity price real exchange rate pass-throughs (Frankel and Saiki (2002) for a fiscal example in Chile related to copper and Edwards (1986) for exchange rate interventions in Columbia related to coffee).

This research on commodity volatility and on the Dutch disease relies on sound and sophisticated econometric techniques. Chapters 1 and 2 investigate short term dynamics of the second moment and are based on time series techniques deviced for capturing the memory in the volatility, as well as the feedback effects of the inventory levels. Asymmetric univariate models such as the GJR-GARCH and exponential GARCH, and multivariate models such as the asymmetric BEKK-GARCH models serve our analysis. Chapters 3 and 4 use tools developed at the crossing of time series and panel techniques. More exactly, we use second generation non-stationary panel tools, which take benefit from the time and cross-section dimensions of our data, tackle the potential non-stationary problems and at the same time control for potential cross-sectional dependence (arising from global factors for example).

As a rabbi once advised Churchill on how to bring about Germany's defeat, there are two possible ways to handle the natural resource curse, one involving natural means, the other supernatural. The natural means would be if a million angels with flaming swords were to descend on the damned resources and destroy it. The supernatural would be to device a sound and balanced policy mix to manage the curse. A realist would choose the natural method, angels with flaming swords. This thesis aims to help those chosing the supernatural one. Our first two Chapters indeed help to better understand the specific volatility dynamics and the way to mitigate the risks by taking into account the inventory tightness feedback. Our last two Chapters show how the Dutch disease may arise from high specialization in the export of one commodity and the structural factors amplifying or limiting it.

Chapter 1

Commodities Volatility and the Theory of Storage

Abstract

The theory of storage has many implications for commodity prices. One of them states that commodity price volatility should increase when inventories are low. This study is the first to investigate systematically the volatility implication of the theory of storage for a large panel of commodity types (agriculturals, metals, precious metals, tree crops and an index). Since inventories are hard to measure and define, especially for high frequency data, we propose to use positive return shocks as a new original proxy. Indeed, positive price shocks potentially indicate a deterioration of the state of the inventories, thereby feed the volatility. We document this volatility feature by estimating asymmetric GARCH models for 17 commodity return series, on the period 1994-2011. We find that this "inventory effect" is relatively robust to different specifications, samples and frequencies. Though not detected on all commodities, we find that this effect is not specific to one type of commodities. We finally illustrate how asymmetric models can be relevant for Value-at-Risk forecasting on commodity price series.

Keywords: Asymmetries * Commodities * GARCH * Inventory * Value-at-Risk.

JEL Classification: C22, G13, Q14.

1.1 Introduction

Pricing theories for commodities have not fundamentally changed over the last decades, excepted some adjustments to the theory of storage. This field now raises a renewed interest for three reasons: (i) a significant commodity price boom occured over the period 2005-2011, commodity indices more than doubled, oil prices were multiplied by three, silver prices by four, etc.; (ii) commodity derivatives occupy a much more important place than before - the Bank of International Settlement thus reports that the amounts outstanding of over-the-counter commodity derivatives (forward and options) exceed 3 trillions dollars as of June 2011; (iii) the pressure on the demand side with the fast development of the BRIC countries, combined with the announced production peak of some commodities raises the questions of the occurence of potential stock-outs.

Understanding commodity prices dynamics is essential for many agents. First, some countries base their economic development on the production and exportation of commodities. Such countries are highly exposed to commodity price fluctuations. An efficient macroeconomic policy would require good instruments and models to predict and manage such price fluctuations. In this context, developing appropriate volatility models is especially important. Second, commodities take an ever-growing share in portfolio strategies. Commodity prices are often found to be negatively correlated with equity and bonds, and to provide in the long run returns close to those offered by equity investments. Since commodities are physical assets with specific constraints (such as the non-negativity of inventories), specific volatility models must be developed for commodity prices to optimize their risk management techniques.

The reference theory for commodity prices and commodity volatility is the theory of storage. This theory explains the difference between spot and futures commodity prices in terms of storage cost, forgone interest and "convenience yield". The convenience yield (or its modern version) is the discounted value related to the potential premiums that can earn a stockholder by selling the stocks on hand in case of a sudden supply disruption or stock-out. As detailed by Ng and Pirrong (1994) in their study of metals price volatility, the theory of storage has some implications that can be relevant to build a model for volatility. One of these implications states that the variance of the commodity price increases in time of low inventories. Indeed, the convenience yield has a stable value, close to zero, when the inventories are abundantly furnished. A decline in the inventories on the contrary causes the convenience yield to increase substantially when stocks are small. The commodity price is thus more volatile in this case. One of our objectives is to document this asymmetry and show how to capture it.

The empirical contributions testing the theory of storage first related the commodity spot futures spread to the state of the inventories (Brennan (1958) and Telser (1958)). Since inventories are not easy to define and measure, later measures relied on various proxies such as monthly dummies for agriculturals (Fama and French (1987)), business cycles for metals (Fama and French (1988)) or, more recently, price shocks (Gorton, Hayashi, and Rouwenhorst (2007)) to highlight the negative dependence of the convenience yield. A second wave of studies then studied the volatility implications of the theory of storage by taking the commodity spot price spread as an indirect measure of the inventory level and looking at its impact on the volatility (Ng and Pirrong (1994), Gao and Wang (2005) and Lien and Yang (2008)). Our point is to test the volatility implication of the theory of storage by relying on positive price shocks.

Our contribution in this field is twofold. First, this study is the first to investigate systematically the volatility implication of the theory of storage for a large panel of commodity types (agriculturals, metals, precious metals, tree crops and an index). Other contributions either focus on some specific commodities (metals for Ng and Pirrong (1994)) or focus on the hedging implications of the theory of storage (Gao and Wang (2005) and Lien and Yang (2008)). Second, this study replaces past spot futures spreads, monthly dummies or business cycles by past positive returns as new proxy for the states of inventories, thereby following the suggestion made by Gorton, Hayashi, and Rouwenhorst (2007). Indeed, declines in inventories can be signalled by positive price shocks. We test the implications of the theory of storage by estimating a GARCH model with an asymmetric term capturing the specific impact of past positive shocks, a model called GJR-GARCH and developed by Glosten, Jagannathan, and Runkle (1993). Since this asymmetry is directly related to the theory of storage, and in particular to the inventories, we label this positive effect on the variance of past positive shocks as the "(commodity) inventory effect", by analogy with the "leverage effect" commonly found on equities (for which past negative, instead of positive, shocks matter).

We work on a set of 17 commodity prices over the period 1994-2011 on a weekly basis and estimate in one-step an ARMA-GJR-GARCH model with Gaussian distribution by maximum likelihood. We then propose for robustness check some alternative specifications, frequencies, samples and distributions. Our results tend to support the theory of storage and in particular the inventory effect. This effect is however not observed for all commodities, and not specific to one type of commodity. It appears that no clear generalizations are possible and that the relevance of asymmetric models has to be examined on a case-by-case basis.

The existence of the commodity inventory effect has some implications for risk management. To illustrate this, we propose in the final section an application where we test whether an appropriate modelling of the commodity prices improves significantly Value-at-Risk (VaR) evaluations. Based on Engle and Manganelli (2004), we find that taking the inventory effect into account improves the VaR estimates for most commodities. A careful investigation of the inventory effect is by consequence relevant before modelling the price of a commodity.

This paper is structured as follows. Section 1.2 presents the theory of storage and its implications for the volatility. Section 1.3 describes the data. Section 1.4 presents the methodology, the GARCH models designed to capture this asymmetry and the results. Extensions including robustness checks and a risk management exercice are then proposed in Section 1.5. Finally, Section 1.6 concludes.

1.2 Theory of Storage

The theory of storage, in its classical version (Kaldor (1939) and Working (1949)), explains the difference between spot and futures commodity prices in terms of storage cost, forgone interest and "convenience yield". Convenience yield is for commodities what dividend yield is for equity stocks. It denotes the stream of implicit benefits that an inventory holder receives by having the commodity on hand (i.e. by the ability to respond more flexibly to unexpected excess demand or supply disruptions). More formally,

$$F_{t,T} - S_t = W_{T-t} + r_t S_t - C_t(I_t) \tag{1.1}$$

where S_t is the spot price of the commodity at date t, $F_{t,T}$ is the price at t of the future contract on the commodity with delivery at time T, W_{T-t} is the global cost of storage between date t and T, r_t the interest rate, $C_t(I_t)$ the convenience yield and I_t the state of the inventories. The difference between $F_{t,T}$ and S_t, also known as the "basis", can be sometimes positive (the market is said to be in contango) or negative (the market is said to be in backwardation). In case of backwardation, the convenience yield gets larger than the storage costs and forgone interests together.

The variation of the convenience yield is the core of the theory of storage and helps to understand how supply and demand fundamentals affect the basis. When the economy builds up inventories, the potential for stock holders to benefit from the stocks on hand decreases. In other words, the convenience yield decreases. By reduction in inventory build-ups on the contrary, stockholders increase their probability to benefit from a sudden positive spike in spot price. The convenience yield in this case increases. Since supply is less elastic in the short run than in the long run, spot prices will be more affected than futures prices. As a result, the basis depends negatively on the state of inventories.

The modernized version of the theory of storage (Deaton and Laroque (1992)) replaces the idea of convenience yield by the probability of occurence of a stock-out. While in the classical theory, an inventory holder benefits from a convenience yield, the modernized version considers that the benefit is related to the implications of a potential stock-out. In both approaches, inventories are the core concept and determine to which extent cost-and-interest-adjusted basis (in other words the basis) deviates from zero.

Many empirical papers support the theory of storage. First empirical studies (Brennan (1958) and Telser (1958)) were based on inventory data and found the negative dependence of the cost-and-interest-adjusted basis (i.e. the convenience yield). This approach is however hard to implement due to the difficulty to define an appropriate scope of relevant inventory data (geographically but also due to product differentiations) and due to the difficulty to collect appropriate data on inventories as they are generally not exhaustively known[1]. As a consequence, a second wave of applied works relied on various proxies for inventories such as monthly dummies for agriculturals (Fama and French (1987)), business cycles for metals (Fama and French (1988)) or, more recently, price shocks (Gorton, Hayashi, and Rouwenhorst (2007)) to highlight the negative dependence of the convenience yield.

Beside these direct tests of the theory of storage, other studies explored some indirect implications of the theory (Ng and Pirrong (1994), Gao and Wang (2005) and Lien and Yang (2008) among others). Ng and Pirrong (1994) list, and test for metals, six implications for the variances and covariance of the spot and futures returns: (1) spot and futures correlation should be positively correlated to the level of the inventory; (2) spot variances should be larger than futures variance; (3) spot and futures variance should be equal when inventories increase; etc. Their results all support the theory of storage. Lien and Yang (2008) test the relevance of including the basis in their variance-covariance specification of spot and futures returns to improve the precision of a dynamic optimal hedge ratio (see Chapter 2 of this thesis for more details). Gao and Wang (2005) finally propose a unified test incorporating aspects of both direct and indirect approaches of the theory of storage.

The estimation framework changes from one contribution to the other, but they share one feature: they all involve both the spot and futures returns. Baillie and Myers (1991) use a bivariate VAR of spot and futures returns and BEKK-GARCH specification for the variance-covariance equations. Ng and Pirrong (1994) use a bivariate ECM-VAR of spot and futures returns where

[1] Improving the transparency and data availability on physical commodity markets is one of the recommendations made by the G20 in 2011.

the correction term is the interest adjusted basis, and include the squared interest-adjusted basis as regressor in the variance equation. Lien and Yang (2008) use the same framework to study optimal futures spot hedge ratios but adjust the variance equation by allowing different effects for negative and positive basis. Gao and Wang (2005) study the interest adjusted basis in a univariate framework with an autoregressive structure in the mean equation and an EGARCH specification in the variance.

We depart from these approaches for different reasons. First, the basis is not the sole proxy for capturing the state of the inventories. As pointed out by Gorton, Hayashi, and Rouwenhorst (2007), the sign of past returns also brings information on the state of the inventories. Positive shocks tend to indicate a deterioration of the stocks. We thus use a "new" (and very simple) proxy which is the sign of the past return, following the suggestion made by Gorton, Hayashi, and Rouwenhorst (2007). Second, our focus is not the relationship between spot and futures returns but the dynamic of the variance of the spot returns. We thus depart from the classical approach related to the theory of storage and focus on spot returns. Third, using the return sign as a proxy could be particularly relevant and useful for commodities where the futures market is not well developed and the basis not available over a sufficient period (steel for example).

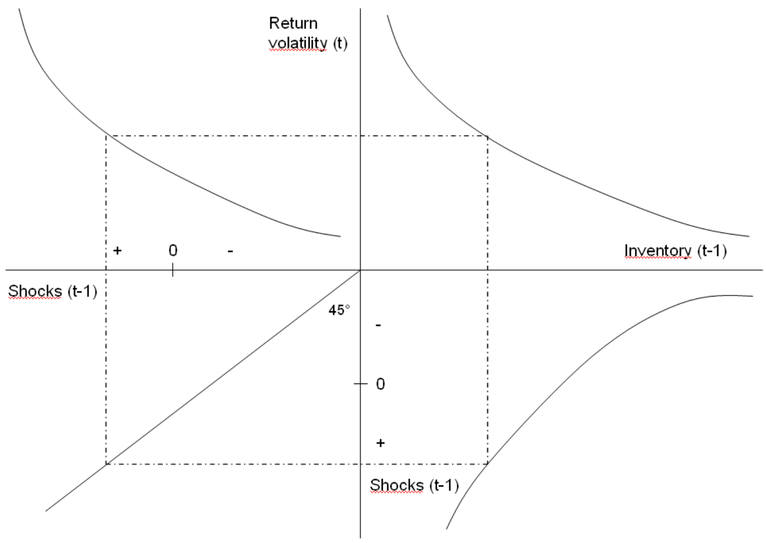


Figure 1.1: Relation between the return volatility, the inventory and the price shocks

As depicted in Figure 1.1 the spot return volatility is expected to be negatively related to lagged inventories (upper right corner of the figure). Since those inventories are not observable, we use as proxy the past shocks, which are expected to be on average positive when inventories are low and negative when invetories are high (lower right corner of the figure). We thus estimate whether positive shocks have a positive effect on the variance (upper left corner of the figure).

1.3 Data

We use a set of 17 commodity spot price series, available in Datastream, which covers the period from 3rd January 1994 to 30th December 2011, 18 years of data on a weekly basis, so 939 observations per series. Our dataset includes the S&P Goldman Sachs Commodity Index[2], five agriculturals (corn, cotton, soybean, sugar and wheat), two tree crops (coffee and rubber), two energy (Brent and gas), four metals (aluminium, copper, nickel and zinc) and three precious metals (gold, palladium and silver). The commodity prices are described in Table 1.1.

We selected these 17 commodity price series on basis of their availability over a sufficiently long period and of their market liquidity. The graph of the series in level and log-returns are presented in Figures 1.2 to 1.10 in the Appendix. We can see in the graphs that most price series have a boom after 2005, followed by a bust for some of them. Due to the higher volatility over the period 2005-2011 we split the sample into two subsamples (1994-2004 and 2005-2011) and reiterate our estimates in the Robustness section.

Our dataset includes very different commodities which follow different dynamics. Supply and demand factors exert their influence on commodities differently. Agriculturals are affected by seasonality and metals by business cycles. Shocks may also have different implications for different commodities. The impact of a supply shock on agricultural prices has in principle no impact the next year. Tree crops (coffee and rubber) on the contrary may have more lasting effects in case of supply shocks (strong winds or rains detrimental to the trees for example). The cost of storage is another parameter of differentiation for commodities. The relative cost of storage is typically lowest for precious metals, low for metals, large for agriculturals and very large for animals or perishable products. In addition, though often considered as a unique asset class, commodities are neither clearly correlated (Erb and Harvey (2006)) nor influenced by the same macroeconomic factors (Batten, Ciner, and Lucey (2010)). As a consequence, we study the commodities separately in the empirical section and interpret the results by taking into account these

[2]The index is a world-production weighted index which comprises 24 commodities from all commodity sectors - energy products, industrial metals, agricultural products, livestock products and precious metals.

Table 1.1: Data description

Series	Description	Type	Unit Base
CORN	Corn No.2 Yellow Cents/Bushel	Agriculturals	Bushel
COTT	Cotton,1 1/16Str Low -Midl,Memph Cts/Lb	Agriculturals	Pound
SOYB	Soyabeans, No.1 Yellow Cts/Bushel	Agriculturals	Bushel
SUGR	Raw Sugar-ISA Daily Price c/lb	Agriculturals	Pound
WHEA	Wheat, No.2 Hard (Kansas) Cts/Bushel	Agriculturals	Bushel
COFF	Coffee-Brazilian (NY) Cents/lb	Tree crops	Pound
RUBB	Rubber (MRE) SMR GP FOB Sen/Kg	Tree crops	Kilogram
BRNT	Crude Oil-Brent Current Month FOB USD/BBL	Energy	Barrel
GAS	Gas Oil-EEC CIF Cargos NWE USD/MT	Energy	Metric Tonne
ALUM	LME-Aluminium 99.7% Cash USD/MT	Metals	Metric Tonne
COPP	LME-Copper, Grade A Cash USD/MT	Metals	Metric Tonne
NICK	LME-Nickel Cash USD/MT	Metals	Metric Tonne
ZINC	LME-SHG Zinc 99.995% Cash USD/MT	Metals	Metric Tonne
GOLD	Gold Bullion LBM USD/Troy ounce	Precious	Troy Ounce
PALD	Palladium USD/Troy Ounce	Precious	Troy Ounce
SILV	Silver Fix LBM Cash Cents/Troy ounce	Precious	Troy Ounce
GSCI	S&P GSCI Commodity Spot Index	Index	Points

Notes. "LME" holds for London Metal Exchange, "MT" for Metric Tonnes, "SHG" for Special High Grade, "LBM" for London Bullion Market, "lb" for pound-mass, "FOB" for Free on Board, "CIF" for Cost, Insurance and freight included, "NWE" for Northwest Europe, "EEC" for European Economic Community, "BBL" for oil barrel, "SMR" for Standard Malaysian Rubber, "MRE" for Malaysian Rubber Exchange, "GP" for General Purpose and "GSCI" for Goldman Sachs Commodity Index.

different features.

We find that all the logarithmics series possess a unit root (see Table 1.6 in the Appendix) and remove the non-stationarity of the logarithmic prices by taking the differences, which are all stationary at 1% according to ADF tests (results available on request). We report in Table 1.2 the descriptive statistics of the log-returns.

Table 1.2: Descriptive statistics for commodity log-returns

Variable	**Mean**	**Std. Dev**	**Min**	**Max**	**SK**	**EK**
DLCORN	0.08%	4.29%	-21.1%	22.3%	-0.403	2.670
DLCOTT	0.03%	4.06%	-14.6%	17.4%	0.192	1.353
DLSOYB	0.06%	3.66%	-24.5%	12.3%	-0.798	3.728
DLSUGR	0.08%	4.51%	-21.8%	18.1%	-0.443	2.031
DLWHEA	0.05%	4.09%	-17.5%	22.9%	0.189	2.605
DLCOFF	0.14%	5.65%	-27.5%	48.5%	0.703	7.723
DLRUBB	0.16%	3.08%	-23.1%	12.0%	-1.006	6.775
DLBREN	0.21%	4.93%	-22.0%	15.1%	-0.592	1.193
DLGAS	0.20%	4.61%	-19.8%	15.4%	-0.475	1.190
DLALUM	0.06%	3.00%	-17.8%	9.2%	-0.405	2.133
DLCOPP	0.16%	3.80%	-25.2%	13.5%	-0.844	4.856
DLNICK	0.14%	5.16%	-22.2%	32.0%	0.069	2.571
DLZINC	0.07%	4.08%	-19.8%	15.9%	-0.333	2.511
DLGOLD	0.15%	2.33%	-13.3%	13.1%	-0.161	4.174
DLPALD	0.17%	4.94%	-24.6%	23.4%	-0.171	3.297
DLSILV	0.18%	4.30%	-35.3%	25.6%	-0.932	8.261
DLGSCI	0.14%	3.14%	-20.6%	12.1%	-0.833	3.763

Notes. 938 weekly observations. All statistics are computed on commodity log-returns. SK stands for Skewness and EK for Excess Kurtosis.

As reported in Table 1.2, the weekly average log-returns are positive for all series. We see that agriculturals average return, from 0.03% to 0.08%, or on an annual basis from 1.5% to 4.1%, are very low compared to other financial assets (5.5% for the S&P 500 over the same period). Other commodities average returns are larger and closer to stock performances, as documented by Erb and Harvey (2006). Regarding the variance, no clear differences appear according to commodity types. Contrary to Gorton and Rouwenhorst (2006), Gorton, Hayashi, and Rouwenhorst (2007), Deaton and Laroque (1992) we find that most of our series exhibit negative skewness (as commonly found for equity stocks). Negative spikes tend to dominate positive spikes in our sample. Without surprise, we also find that all series (espacially aluminium, coffee and rubber) are leptokurtic. To take into account the potential effects of the non-zero skewness and excess kurtosis, we reiterate our estimates with a skewed-student distribution (as the one proposed by Lambert and Laurent (2002)) in the Robustness section. We use log-returns in the empirical sections.

1.4 Methodology and Results

GARCH models were initiated by Engle (1982) and Bollerslev (1986) in the eighties to capture the time varying volatility of financial series. The models were extended in the nineties so that we now face a genuine "ARCH-mada" of models (see Terasvirta (2006), Bollerslev (2008) and Bauwens, Hafner, and Laurent (2011) for some review in a univariate framework; Bauwens, Laurent, and Rombouts (2006) and Silvennoinen and Terasvirta (2008) in the multivariate one). Combined with an ARMA conditional mean, GARCH family models are specified as follows:

$$y_t = \delta + \phi_1 y_{t-1} + \phi_2 \epsilon_{t-1} + \epsilon_t \tag{1.2}$$

$$\epsilon_t = z_t \sigma_t \tag{1.3}$$

$$z_t \rightarrow i.i.d. \quad D(0,1) \tag{1.4}$$

$$\sigma_t^2 = E(\epsilon t^2|\Omega_{t-1}) \tag{1.5}$$

where y_t is the raw series, ϵ_t is the error term, σ_t^2 is the conditional variance, z_t is the standardized error, identically and independently distributed with mean zero and unit variance, and Ω_{t-1} is the information set at time t-1. GARCH models thus differ in the way that σ_t^2 is specified. The most common specification is the GARCH(1,1) where $\sigma_t^2 = \omega + \alpha\epsilon_{t-1}^2 + \beta\sigma_{t-1}^2$.

Among the GARCH models, some are developed to cope with the characteristic, originally put forward by Black (1976), that the increase in volatility of equity stocks is usually larger when the returns are negative than when they are positive. The traditional explanation of this so-called "leverage" effect is related to the financial structure of equity stocks. Falling returns give rise to a deterioration of the debt-to-equity ratio, which raises the probability of a default. Some common models capturing this asymmetry are the GJR-GARCH of Glosten, Jagannathan, and Runkle (1993) and the EGARCH of Nelson (1991), among others. They respectively specify the conditional variance as follows:

$$\sigma_t^2 = \omega + \alpha\epsilon_{t-1}^2 + \gamma S_{t-1}^{-}\epsilon_{t-1}^2 + \beta\sigma_{t-1}^2 \tag{1.6}$$

$$log(\sigma_t) = \omega + \beta log(\sigma_{t-1}) + \theta_1\epsilon_{t-1} + \theta_2(|\epsilon_{t-1}| - E|\epsilon_{t-1}|) \tag{1.7}$$

where S_{t-i}^{-} is a dummy variable with value 1 if ϵ_{t-1} is negative and 0 otherwise, γ is the parameter capturing the asymmetric effect in the GJR-GARCH model, θ_1 and θ_2 are the parameters capturing respectively the sign effect and the amplitude effect of past shocks in the EGARCH model. A positive γ, or a negative θ_1, means that a past negative return has a larger impact on conditional volatility than a past positive return of the same amplitude. We use the GJR-GARCH(1,1) and EGARCH(1,1) models in our empirical sections to capture the commodity inventory effect. Since positive shocks have a larger effect on the conditional variance than negative shocks, according to the commodity inventory effect, we expect γ, the asymmetric term of the GJR-GARCH model, to be negative and θ_1, the asymmetric term of the EGARCH model, to be positive.

We estimate ARMA(1,1)-GJR-GARCH(1,1) models by maximum likelihood in a one-step procedure, based on Gaussian distribution. We report in Table 1.3 the value of the asymmetry coefficient and its p-value. A negative sign for γ confirms the commodity inventory effect. In other words, lower inventories, proxied by positive returns, tend to increase the conditional variance, as the theory predicts.

We first note that the γs, or asymmetry coefficients, of 15 series (out of 17) have a negative sign, as expected. Second, six are significant at 10%. Third, we check by looking at Akaike Information Criterion whether adding the asymmetry term brings relevant information. Therefore, we compare the AIC for GJR-GARCH(1,1) and GARCH(1,1) models and find that the former is preferred for 9 commodities.

Regarding agriculturals, sugar is the sole commodity for which an inventory effect is significant. We find significant effects for other agriculturals with alternative specifications, but not in the benchmark case. We could expect stronger results for agriculturals. As stated by Fama and French (1987), commodities with high cost of storage (such as agriculturals) should have large inventory effect. Indeed, in case of high storage cost, the inventory buffer is expected to be lower. In such as case, a decline in inventory is expected to have a direct and strong effect. For precious metals as well, the idea that the cost of storage is determinant for the inventory effect is refuted since a strong effect is found for two precious metals, contrary to what Ng and Pirrong (1994) found for silver. Indeed, precious metals have low storage cost and should have accordingly a weaker inventory effect.

These results globally confirm that there is an inventory effect, but also that this effect is not effective for all series. It appears that no generalizations are possible and that the relevance of asymmetric models has to be examined on a case-by-case basis. It also appears that cost of storage is not the core determinant of the inventory effect. Other parameters should play a role, especially the short term supply elasticity. Huge stocks of gold do not mean that gold holders are willing to sell it directly in case of supply/demand shocks. This view tends to

Table 1.3: Commodity Inventory Effect, captured by the asymmetric term γ of an ARMA-GARCH-GJR(1,1) estimated on weekly data over the period 1994-2011

		GJR(1,1)-Gaussian		Akaike Information Criterion		
	Robustness	γ	p-value	GJR	GARCH	GJR pref
CORN	0	-0.042	0.46	-3.6197	-3.6212	No
COTT	0	-0.008	0.77	-3.6746	-3.6788	No
SOYB	1	-0.032	0.49	-3.9509	-3.9525	No
SUGR	3	-0.025	**0.05**	-3.4189	-3.4200	No
WHEA	1	-0.009	0.85	-3.6748	-3.6769	No
COFF	4	-0.171	**0.02**	-3.0355	-3.0172	Yes
RUBB	2	-0.076	0.40	-4.4617	-4.4624	No
BRNT	0	0.070	0.24	-3.2235	-3.2214	Yes
GAS	2	-0.053	0.11	-3.3615	-3.3605	Yes
ALUM	2	-0.022	0.42	-4.3056	-4.3047	Yes
COPP	0	0.036	0.42	-3.8750	-3.8755	No
NICK	0	-0.016	0.57	-3.2194	-3.2211	No
ZINC	4	-0.078	**0.00**	-3.8733	-3.8629	Yes
GOLD	4	-0.161	**0.02**	-5.0126	-4.9979	Yes
PALD	1	-0.057	0.19	-3.3634	-3.3629	Yes
SILV	5	-0.115	**0.00**	-3.7355	-3.7071	Yes
GSCI	2	-0.051	**0.10**	-4.2253	-4.2226	Yes

Notes. Weekly data. Column 2 "Robustness" indicates the number of specifications where the commodity inventory effect is significant at 10%. These specifications are reported in the Robustness section. For example, gold log returns have a commodity inventory effect significant at 10% in 5 specifications. Columns 3 and 4 report the estimates and p-value of the asymmetric term γ from an ARMA-GARCH-GJR(1,1) specification with Gaussian distribution and constants in the mean and variance equations. Columns 5, 6 and 7 report Akaike Information criteria (to be minimized) for GARCH-GJR(1,1) and GARCH(1,1) specifications, respectively, and whether the GJR must be preferred.

give more support to the classical version of the theory of storage than to the modern one. The modern version considers that there is no convenience yield but a probability of stock-out that could give to stockholders an opportunity to set higher prices. Such probability of stock out is certainly not relevant for gold, since the annual demand oscillates around 4,000 tons compared to above-ground stocks of 158,000 tons. We test in the next section if these results hold on alternative specifications, samples, distributions and frequencies.

1.5 Extensions

We propose two extensions in this section. The first one explores the robustness of our results, by modifying various parameters of our benchmark approach. The second one illustrates the relevance of considering the inventory effect for risk management.

1.5.1 Robustness checks

We report in Table 1.4 the results of five alternative specifications. First, the benchmark estimates rely on weekly data. Working on weekly data seems an appropriate balance between the objective of reducing the daily noise in the data and the objective of keeping a sufficient sample size. We check in this extension if the results hold for daily data. Second, most graphs depict an increase in volatility from around 2005, what some called the third commodity boom (Radetzki (2006)). We thus reiterate our estimates on a subsample of 573 observations over the period 1994-2004 and another subsample of 365 observations over the period 2005-2011. Then, since we found negative skewness for most series and non-zero excess kurtosis for nearly all series, we reiterate our estimates by replacing the Gaussian by a skewed-student distribution (as the one proposed by Lambert and Laurent (2002)). Finally, we replace the GJR-GARCH by the EGARCH model, an alternative asymmetric GARCH model developed by Nelson (1991) where the parameter of interest is now θ_1.

Some of these alternative specifications give even stronger results for the inventory effect (daily frequency model and first subsample model), others weaker results (the second subsample model and the EGARCH model). As a whole, these alternative results tend to confirm the benchmark results.

Regarding the estimates on the daily observations, we find that now three agriculturals have a significant (at 10%) inventory effect (wheat, corn and soybean). This confirms the results of Benavides-Perales (2010) on corn and wheat daily data as well, where they use the cost-adjusted basis as proxy for the inventories. Regarding metals, aluminium and zinc have now a significant (at 10%) inventory effect, as well as gold and silver. Ng

Table 1.4: Commodity Inventory Effect captured on alternative specifications, frequencies, samples and distributions

Model	GJR		GJR		GJR		GJR		EGARCH	
Frequ.	Daily		Weekly		Weekly		Weekly		Weekly	
Sample	1994-2011		1994-2004		1995-2011		1994-2011		1994-2011	
Distrib.	Gaussian		Gaussian		Gaussian		Sk.-Student		Gaussian	
	γ	p-val	γ	p-val	γ	p-val	γ	p-val	θ_1	p-val
CORN	0.008	0.06	-0.024	0.76	0.052	0.62	0.019	0.81	0.018	0.50
COTT	0.003	0.69	0.021	0.63	-0.045	0.37	-0.014	0.63	-0.010	0.44
SOYB	-0.048	0.00	0.004	0.95	-0.042	0.67	-0.048	0.34	0.015	0.54
SUGR	0.002	0.78	-0.055	0.01	-0.152	0.11	-0.011	0.71	0.083	0.09
WHEA	-0.021	0.06	-0.089	0.27	0.023	0.74	-0.015	0.68	-0.016	0.64
COFF	-0.054	0.00	-0.220	0.00	-0.106	0.29	-0.172	0.01	0.051	0.65
RUBB	-0.008	0.74	-0.165	0.01	0.209	0.38	-0.116	0.08	0.030	0.55
BRNT	0.021	0.05	-0.009	0.73	0.109	0.12	0.061	0.11	-0.095	0.15
GAS	-0.003	0.80	-0.114	0.01	0.008	0.87	-0.041	0.09	-0.001	0.97
ALUM	-0.017	0.09	-0.104	0.07	-0.002	0.97	-0.022	0.42	0.006	0.61
COPP	0.013	0.30	-0.029	0.76	0.126	0.11	-0.005	0.87	-0.044	0.19
NICK	-0.003	0.79	-0.032	0.44	0.003	0.95	-0.010	0.70	0.008	0.66
ZINC	-0.022	0.03	-0.148	0.00	-0.007	0.90	-0.073	0.00	0.024	0.13
GOLD	-0.056	0.00	-0.158	0.24	-0.206	0.00	-0.115	0.04	0.061	0.15
PALD	-0.008	0.70	-0.052	0.41	-0.070	0.19	-0.099	0.08	0.068	0.13
SILV	-0.034	0.01	-0.120	0.06	-0.196	0.01	-0.108	0.00	0.040	0.35
GSCI	-0.078	0.42	-0.093	0.09	-0.013	0.88	-0.029	0.19	0.009	0.82
NOBS	4694		573		365		938		938	

and Pirrong (1994) also found on their data a confirmation of the theory of storage for their non-precious metals (including copper), but no effect on silver.

Regarding the estimates on the two subsamples, we note that the results are confirmed on the 1994-2004 period but weaker on the 2005-2011 period. Lower significance in the 2005-2011 subsample is partially due to the lower sample size. In addition, the period is characterized by a high volatility with a boom and bust pattern for many commodities, combined with a high level of macroeconomic uncertainty related to the sub-prime and euro crises. This noise probably also contributed to weaken the results.

Finally, we replaced the Gaussian distribution by a skewed-student distribution. Taking the skewness into consideration does not neutralize the asymmetric inventory effect but, on the contrary, even strenghtens the effects (all precious metals have a significant inventory effect under this specification).[3] The last column reports the estimates from the EGARCH model of Nelson (1991). The signs of the θ_1s are positive for most series, as the theory predicts, but significance is lower than in GJR-GARCH models. The EGARCH specification splits the asymmetric effect into two coefficients, the sign and the amplitude, which might reduce the significance of each parameter examined in isolation.

As a conclusion, we consider that there is an inventory effect, but not effective for all commodities, and not specific to one type of commodity. It appears that no generalizations are possible and that the relevance of asymmetric models has to be examined on a case-by-case basis.

1.5.2 Value-at-Risk application

Accounting for the volatility asymmetry documented in the precedent sections can help to improve risk management and VaR in particular.

Trading in commodities, as an alternative investment class to traditional portfolios comprising equity stocks and bonds, has considerably increased in recent years. Gorton and Rouwenhorst (2006) study the properties of commodity futures by building an equally-weighted index of commodity series and find that commodity futures have historically offered the same return and Sharpe ratio as equities. Since return goes hand in hand with risk, we propose in this section an extension to show the relevance of asymmetric models in the context of risk management.

[3] We found similar results with a skewed-student distribution on daily data. We also find in this case a significant inventory effect for palladium. Results non reported, available on request.

Instead of restricting our analysis to a portfolio comprising all our commodities as in Gorton and Rouwenhorst (2006), we follow Erb and Harvey (2006) and Batten, Ciner, and Lucey (2010) who recommand to study individual commodity price series rather than indexes, as they find that commodity futures returns are neither significantly correlated, nor affected by the same macroeconomic factors. We thus consider a risk management exercice on each of our commodities and investigate whether asymmetric GARCH are valid models for VaR forecasting.

A VaR measure can be defined as a quantitative tool whose goal is to assess the potential loss that can be incurred with a certain level of confidence under normal market conditions by an investor over a given time period and for a given investment. In mathematical terms, the VaR is defined to be the α quantile of the investment's profit and loss distribution

$$VaR_t(\alpha) = F^{-1}(\alpha|\Omega_t) \tag{1.8}$$

where $F^{-1}(\cdot|\Omega_t)$ refers to the quantile function of profit and losses distributions which varies over time as market conditions change. Modelling the quantile function requires a model that best captures the dynamic of the conditional mean and variance.

Determining the accuracy of a VaR model can be reduced to checking if two properties are satisfied (see Campbell (2005) and Hurlin and Tokpavi (2007) for recent reviews of backtesting procedures). First, the probability of realizing a loss in excess of the reported $VaR_t(\alpha)$ must be precisely equal to α X 100% (what is called the unconditional coverage property). Second, the cases where the losses exceed the reported $VaR_t(\alpha)$ must be independently distributed (what is called the independence property).

We report in Table 1.5 the results of the Dynamic Quantile (DQ) test, developed by Engle and Manganelli (2004), which tests both properties at the same time, by regressing the $Hit_t(\alpha)$ variable (where $Hit_t(\alpha) = 1 - \alpha$ when the loss exceeds the VaR, and $Hit_t(\alpha) = -\alpha$ otherwise) on a constant (equal to 0 if the unconditional coverage property holds) and on its lagged values (non-significant if the independence assumption holds) and by testing the joint significance of the parameters by a Wald test.

The statistics and p-values of the Engle and Manganelli (2004) DQ test are reported in Table 1.5. We report results for the series where the commodity inventory effect was found to be significant at 10% in the benchmark case (Table 1.3) and for the series which have a significant 10% in at least two of the 5 alternative cases analyzed in the Robustness section (Table 1.4). Since the commodity inventory effect relates positive shocks to the volatility, we focus on the right side of the return distribution by examining the VaR for short trading positions only (Giot and Laurent (2003)).

Table 1.5: Value-at-Risk application: validation of GJR-GARCH(1,1) models by dynamic quantile tests

	γ	Robustness	DQ test Q95		DQ test Q99	
			stat.	p-val	stat.	p-val
SUGR	-0.025*	3	11.678	0.07	2.118	0.91
COFF	-0.171*	4	9.463	0.15	4.317	0.63
RUBB	-0.076	2	7.344	0.29	5.535	0.48
GAS	-0.053	2	9.391	0.15	1.416	0.96
ALUM	-0.022	2	3.557	0.74	11.775	0.07
ZINC	-0.078*	4	3.6282	0.73	0.939	0.99
GOLD	-0.161*	4	3.948	0.68	8.011	0.24
SILV	-0.115*	5	6.113	0.41	2.702	0.85
GSCI	-0.051*	2	3.611	0.73	70.350	0.00

Notes. Weekly data. Column 2 "γ" reports the γ estimates capturing the effect of positive shocks on the variance, as obtained in the benchmark case (cf Table 1.3). Column 3 "Robustness" indicates the number of specifications where the commodity inventory effect is significant at 10%. These specifications are reported in the Robustness section (cf Table 1.4). Columns 4-7 report the Dynamic Quantile test statistics and p-values of Engle and Manganelli (2004) for quantiles 95% and 99%.

We find that the validity of GJR-GARCH(1,1) VaR(95%) and VaR(99%) models cannot be rejected for coffee, rubber, gas, zinc, gold and silver. These results confirm those of Giot and Laurent (2003) who compare different GARCH models performances and find that the one allowing asymmetries (APARCH) performs the best[4]. The validity of the GJR-GARCH(1,1) model is rejected for three cases (sugar for the DQ test Q95, and GSCI and aluminium for the DQ test Q99). Bearing in mind these three exceptions, these results overall confirm the relevance of considering an asymmetric volatility model to capture the dynamics of commodity returns in a risk management context.

1.6 Conclusions

The theory implicitly states that commodity prices volatility should increase when the state of the stocks decline. We test empirically the relevance of

[4] They compare their models according to p-values derived from the Kupiec (1995) test. Since the later only tests the unconditional coverage property and not the independence property, we opted in our paper for the Engle and Manganelli (2004) test.

this statement by using a "new" (and very simple) proxy which is the sign of the past return. Indeed, as suggested by Gorton, Hayashi, and Rouwenhorst (2007), past positive price shocks bring information on the state of the inventories. Positive shocks tend to indicate a deterioration of the stocks. We thus test what we call the "inventory effect" by estimating GARCH models with an asymmetric term capturing past positive shocks (GJR-GARCH model), on a set of 17 commodities (agriculturals, tree crops, metals, precious metals, index) over the period 1994-2011.

As a parallel to the leverage effect requiring asymmetric volatility models for equity, we find that there is also an inventory effect requiring asymmetric volatility models for commodities, that this effect is not detected for all commodities, and not specific to one type of commodity. It appears that the relevance of asymmetric models has to be examined on a case-by-case basis. Our results also note that the inventory effect is strong for precious metals, contrary to the assumption that commodities with low storage cost (like gold and silver) are less affected by the theory of storage. Our interpretation is that the cost is not the sole determinant of inventory supply, short term supply elasticity also matters.

As future areas of research, we consider that the theory of storage does not say enough about the role of inventories and its determinants. Better understanding the microeconomics of the inventory effect would be highly relevant. In addition, the co-dependence of commodity volatility and the dependence on common factors would be a promising area of research. Finally, the inventory effect has a role to play in risk management models, including optimal hedge ratios determination. This is precisely the topic of our next Chapter.

Appendix

Table 1.6: Descriptive statistics for commodity prices in logarithms

Variable	Mean	Std. Dev	Min	Max	ADF	Trend	Lags	Type
LCORN	5.61	0.38	4.98	6.66	-0.98	0	0	Agric.
LCOTT	4.11	0.33	3.28	5.34	-1.72	0	0	Agric.
LSOYB	6.53	0.33	5.97	7.39	-2.57	1	8	Agric.
LSUGR	2.39	0.41	1.56	3.44	-1.19	0	0	Agric.
LWHEA	6.10	0.34	5.48	7.14	-1.74	0	0	Agric.
LCOFF	4.67	0.47	3.62	5.77	-1.50	0	9	Tree crops
LRUBB	6.10	0.55	5.27	7.46	-2.47	1	8	Tree crops
LBRNT	3.56	0.68	2.28	4.98	-2.80	1	0	Energy
LGAS	5.75	0.67	4.54	7.21	-2.55	1	0	Energy
LALUM	7.47	0.25	7.04	8.09	-1.99	0	0	Metals
LCOPP	8.04	0.62	7.20	9.22	-0.98	0	8	Metals
LNICK	9.31	0.61	8.24	10.90	-1.50	0	7	Metals
LZINC	7.20	0.44	6.60	8.43	-1.42	0	0	Metals
LGOLD	6.17	0.53	5.54	7.54	-1.18	1	5	Precious
LPALD	5.67	0.54	4.74	6.99	-1.63	0	6	Precious
LSILV	6.66	0.60	6.01	8.49	-1.79	1	5	Precious
LGSCI	5.66	0.49	4.87	6.79	-2.71	1	8	Index

Notes. 939 weekly observations (1994-2011). All statistics are computed on commodity prices in logarithms. ADF stands for Augmented Dickey-Fuller. The critical values at 1%, 5% and 10% are -3.43, -2.86 and -2.57, respectively, for the ADF t-test without deterministic trend, and -3.96, -3.41 and -3.12 for the ADF t-test with trend. The null of unit root is never rejected in our sample. "Trend" takes the value 1 if a deterministic trend had to be included in the ADF test, and 0 otherwise. "Lags" mentions the number of autoregressive lags that had to be included in the ADF specifications according to the Akaike Information Criterion.

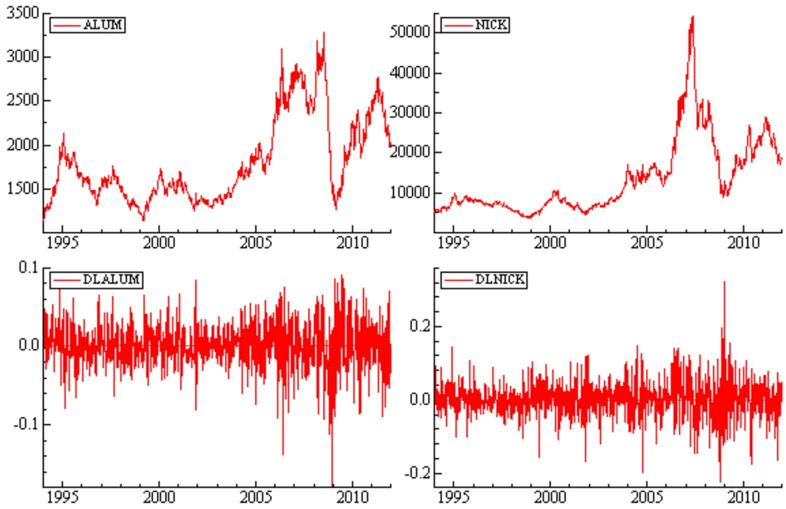


Figure 1.2: Price and log-return series of aluminium and nickel

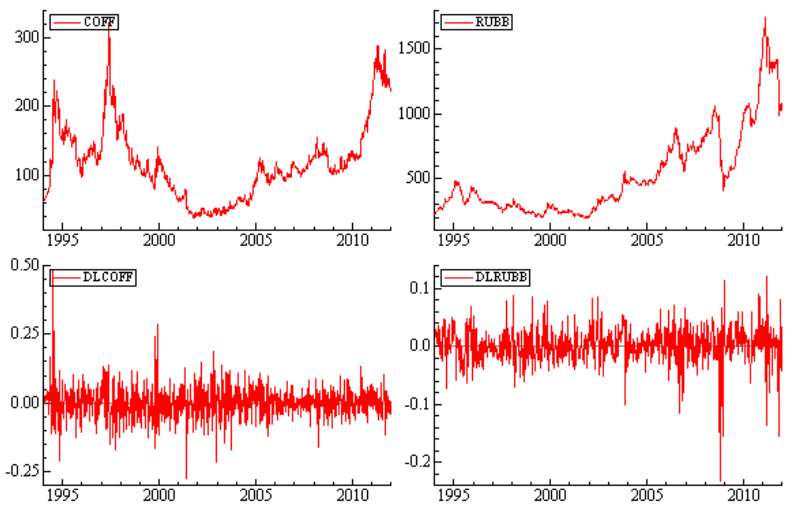


Figure 1.3: Price and log-return series of coffee and rubber

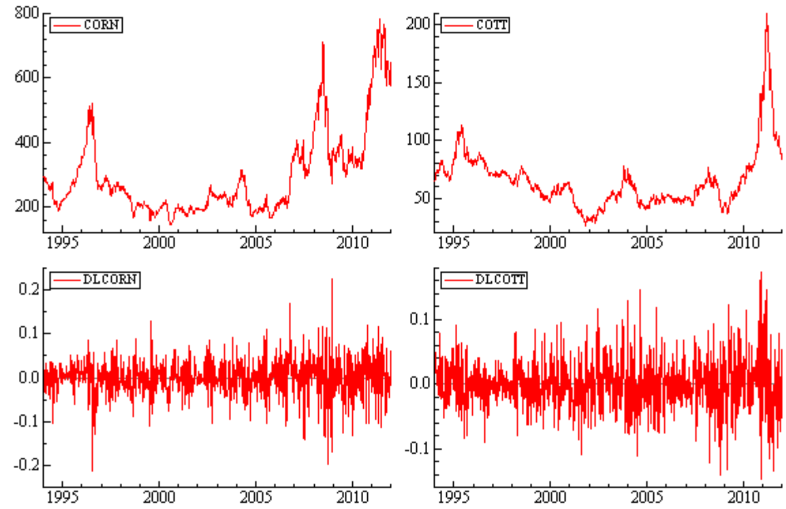


Figure 1.4: Price and log-return series of corn and cotton

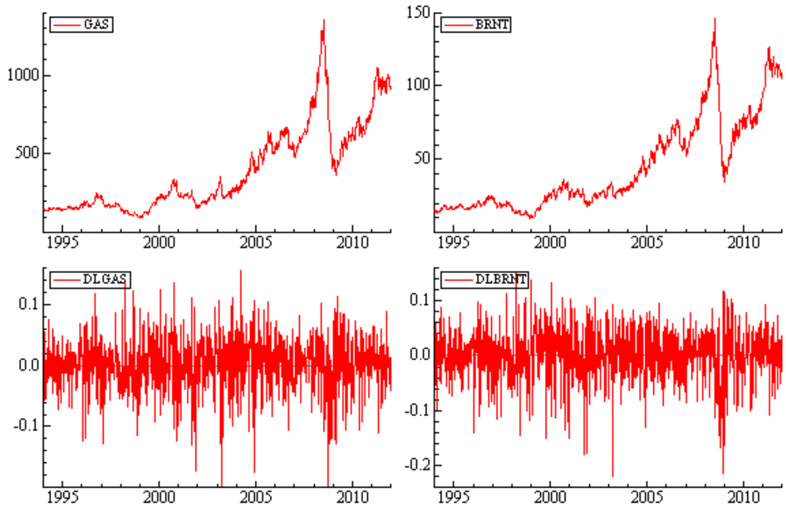


Figure 1.5: Price and log-return series of gas and brent

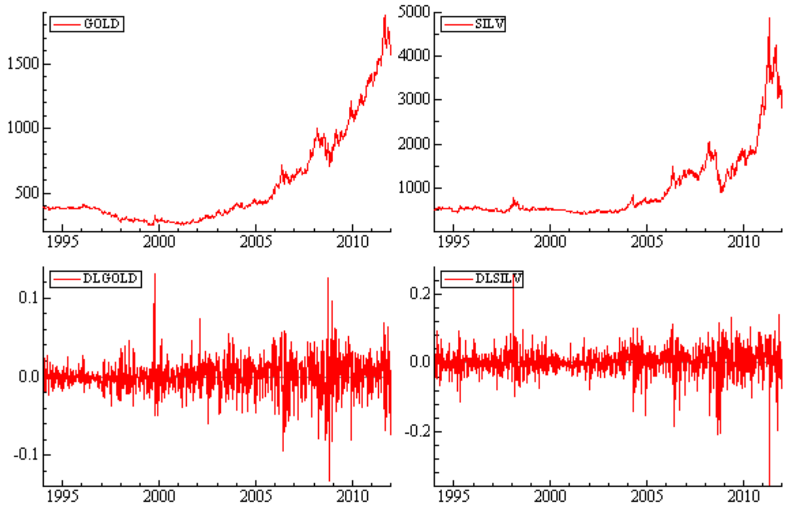


Figure 1.6: Price and log-return series of gold and silver

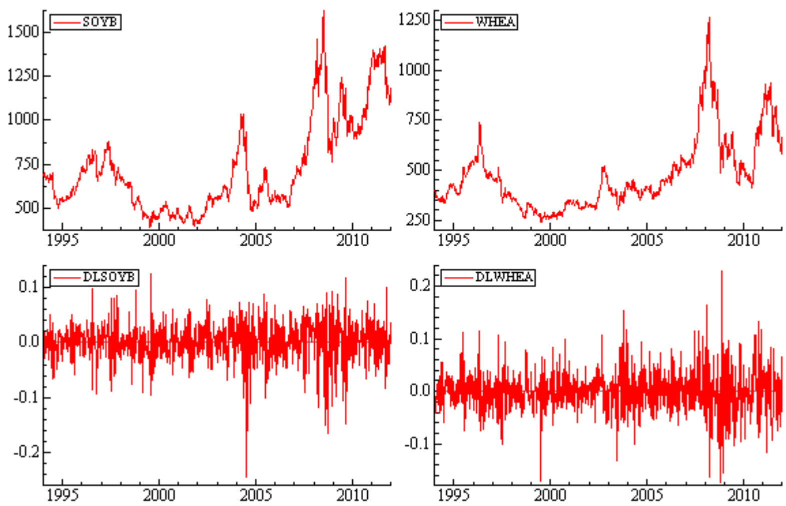


Figure 1.7: Price and log-return series of soybean and wheat

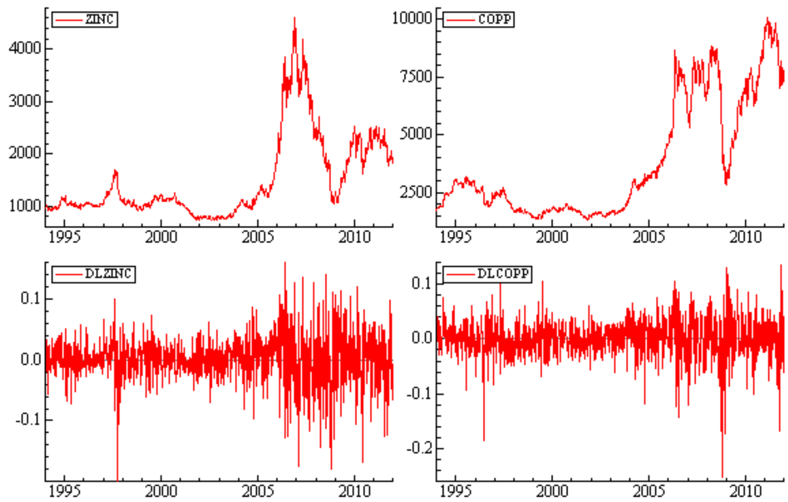


Figure 1.8: Price and log-return series of zinc and copper

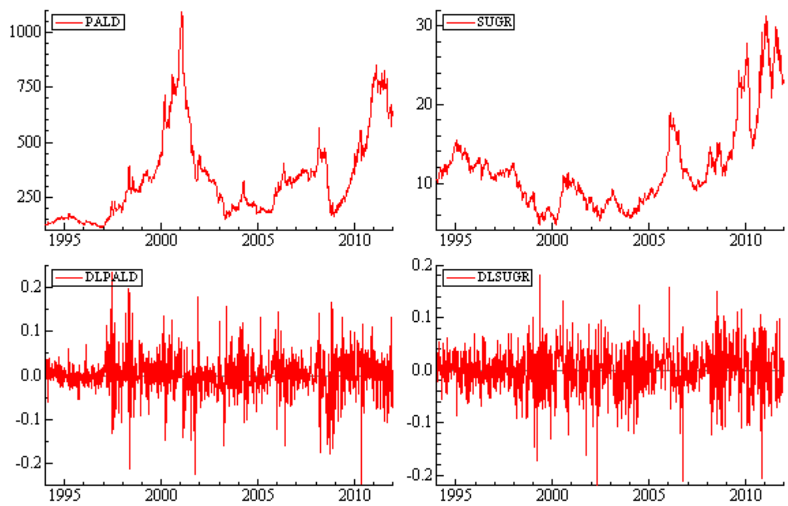


Figure 1.9: Price and log-return series of palladium and sugar

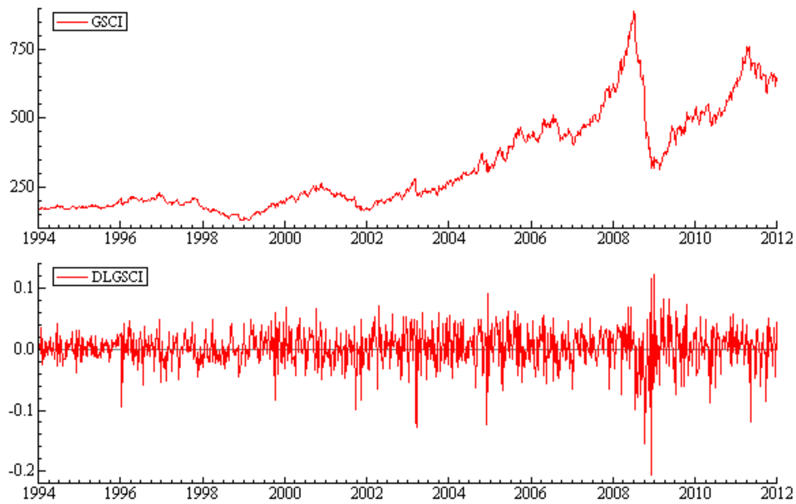


Figure 1.10: Price and log-return series of GSCI Commodity Index

Chapter 2

The Asymmetric Commodity Inventory Effect on the Optimal Hedge Ratio[1]

Abstract

Hedging strategies for commodity prices largely rely on dynamic models to compute optimal hedge ratios. This paper illustrates the importance of considering the commodity *inventory effect* (effect by which the commodity price volatility increases more after a positive shock than after a negative shock of the same magnitude) in modelling the variance-covariance dynamics. We show by in-sample and out-of-sample forecasts that a commodity price index portfolio optimized by asymmetric BEKK-GARCH models outperforms the symmetric BEKK-GARCH, static (OLS) or *naïve* models. Robustness checks on a set of commodities and by an alternative mean-variance optimization framework confirm the relevance of taking into account the inventory effect in commodity hedging strategies.

Keywords: BEKK * Commodity * Asymmetries * Hedging * Inventory effect.

JEL Classification: G13, C32, Q02.

[1] Joint work with Besik Samkharadze. Forthcoming in the Journal of Futures Markets.

2.1 Introduction

Minimizing the investment risk through futures hedging is a technique widely used in many fields (currencies, stocks, commodities). Static regression techniques were first developed to estimate optimal hedge ratios (hereafter referred to as OHR). Dynamic versions, computed with GARCH models, were then proposed and shown to perform better than static ones (Baillie and Myers (1991), Kroner and Sultan (1993) and Lien, Tse, and Tsui (2002) among others).

One shortcoming of these papers is that they overlook the potential asymmetries characterizing the dynamics of some assets' return volatility. This is the case of stock prices, for which the increase in volatility is usually larger when the returns are negative than when they are positive. The traditional explanation of this asymmetry is related to the balance-sheet effect. Falling returns give rise to a deterioration of the debt-to-equity ratio, which raises the probability of default. This effect, originally put forward by Black (1976), is usually referred to as the *"leverage effect"*.

Commodity prices are another example of assets characterized by an asymmetric dynamics in the volatility. On the contrary to the leverage effect, positive shocks to commodity returns give rise to an increase in volatility larger than the one subsequent to a negative shock. This phenomenon was already mentioned by Ng and Pirrong (1994) who, building on the theory of storage, find for metal prices that inventory conditions, proxied by the basis (i.e. the spot-futures spread), affect the volatility of the series asymmetrically. Intuitively, a shortage drives the prices up and increase the volatility/nervosity of the markets, while large inventories, tend to put downward pressure on prices, and lead to relatively low volatility. This inventory effect is documented in Chapter 1 and found to be relatively robust, for different subsamples, for diverse types of commodities and for different ways of specifying the asymmetry, though not so strong as the leverage effect regarding equity stocks.

Some papers take into account asymmetries in their dynamic OHR applications such as Lien and Yang (2006) for currency markets and Brooks, Henry, and Persand (2002) for stock returns. With respect to commodities, Lien and Yang (2008) use a bivariate dynamic conditional correlation model (Engle (2002)) capturing the inventory asymmetries by adding a decomposition of the basis (into negative and positive terms) in the variance equation. Their approach relies on the theory of storage as laid out by Ng and Pirrong (1994) and takes the basis (as decomposed) as proxy of the state of the inventories.

We propose an alternative approach. The basis is not the sole proxy for inventories. As studied by Gorton, Hayashi, and Rouwenhorst (2007), spot return are also good indicators of the state of inventories. In addition, futures data is potentially more affected by speculation related to time-varying risk premium considerations. Since asymmetries are strictly related to inventory data, spot returns are potentially less noisy and good alternatives to the basis. We thus use a model close to the asymmetric BEKK-GARCH specification developed by Kroner and Ng (1998), which is a multivariate extension of the univariate GJR-GARCH model developed by Glosten, Jagannathan, and Runkle (1993). They share the feature of allowing asymmetric shocks in the models, but contrary to their specifications our model (hereinafter Asym-BEKK) captures specifically the positive shocks, instead of the negative ones, in view to highlight commodities inventory effects.

The goals of this paper are twofold. First, we investigate and illustrate the relevance of taking into account the commodities inventory effect in hedging techniques through an Asym-BEKK model. Second, we examine and compare the performance of our methodology through portfolio variance minimization, in-sample and out-of-sample forecasts. This paper is laid out as follows: Section 2.2 introduces to hedging and derives the OHR, Section 2.3 presents the econometric methodology, while Section 2.4 describes the data. Section 2.5 presents the results, Section 2.6 some extensions and Section 2.7 concludes.

2.2 Hedging

A hedge is achieved by taking opposite positions in the spot and futures market at the same time, so that losses supported in one market are, to some extent, offset by opposite price movements in the other market. The number of futures contract units needed to be issued depends on the covariance of the spot and futures contracts, as well as the variances of futures and spot contracts. The OHR of futures to spot contract, the one minimizing the variance of the hedged portfolio, is derived as follows.

Let S_t and F_t denote the logarithm of the spot and futures commodity prices, respectively. The return on a position in spot is $\Delta S_t = S_t - S_{t-1}$ and on a futures position is $\Delta F_t = F_t - F_{t-1}$. The expected return on a portfolio, R_t comprising one unit of the commodity and β units of the futures contract would then be:

$$E_{t-1}(R_t) = E_{t-1}(\Delta S_t) - \beta_{t-1} E_{t-1}(\Delta F_t) \tag{2.1}$$

where β_{t-1} is the hedge ratio for time t determined at $t-1$. The variance of the portfolio is:

$$VAR(R_t) = VAR(\Delta S_t) + \beta_{t-1}^2 VAR(\Delta F_t) - 2\beta_{t-1} COV(\Delta S_t, \Delta F_t) \tag{2.2}$$

Combining the portfolio variance in a two-moment utility function we get:

$$U(E_{t-1}(R_t), VAR(R_t)) = E_{t-1}(R_t) - \theta VAR(R_t) \tag{2.3}$$

where θ is the risk aversion parameter. Maximizing the utility $U(E_{t-1}(R_t), VAR(R_t))$ is equivalent to solve:

$$\begin{aligned} max \quad & E_{t-1}(\Delta S_t) - \beta_{t-1}E_{t-1}(\Delta F_t) \\ & -\theta(VAR(\Delta S_t) + \beta_{t-1}^2 VAR(\Delta F_t) - 2\beta_{t-1}COV(\Delta S_t, \Delta F_t)) \end{aligned} \quad (2.4)$$

Solving Equation 2.4 under the assumption that F_t follows a martingale process, i.e. $E_{t-1}(\Delta F_t) = 0$, yields the OHR β_{t-1}^*:

$$\beta_{t-1}^* = \frac{COV(\Delta S_t, \Delta F_t)}{VAR(\Delta F_t)} \quad (2.5)$$

$$= \rho_t \sqrt{VAR(\Delta S_t)/VAR(\Delta F_t)} \quad (2.6)$$

where ρ_t is the correlation between spot and futures returns.

2.3 Methodology

The simplest way to estimate the OHR laid out in Equation 2.6 is to assume that spot and futures returns are perfectly correlated and the spot and futures variances equal to each other. Under these assumptions, the OHR is $\beta_{t-1}^* = 1$. This is the *naïve* model, that we use later as comparison point for more sophisticated models.

Another simplist approach is to assume that the variances and covariance of spot and futures returns are constant over time. In this case, $\beta^* = \frac{COV(\Delta S, \Delta F)}{VAR(\Delta F)}$ is just the coefficient of a linear regression of the spot return on the futures return. This is the OLS model, that we also use later as a benchmark.

As reported by Baillie and Myers (1991), Kroner and Sultan (1993) and Lien, Tse, and Tsui (2002), early models assuming that the variance-covariance matrix of returns is constant over time proved too restrictive. Multivariate GARCH models relax such assumption and are largely used in this context (see Bauwens, Laurent, and Rombouts (2006) for a survey of MGARCH models).

Kroner and Sultan (1993) use the constant conditional correlation (CCC) model of Bollerslev (1990). The main weakness of this model is that conditional variances may vary over time but not covariances. As documented in Ng and Pirrong (1994), the covariance of spot and futures commodity returns tend to decrease when the inventories get tight, according to the theory of storage. Therefore, this multivariate GARCH model is not well-suited for our purposes.

Lien, Tse, and Tsui (2002) use a dynamic version of the CCC called dynamic conditional correlation model, as developed by Engle (2002), where the correlation is allowed to vary over time. Again, this model is rather restrictive in the sense that the correlation is only affected by its own past and cannot be affected, for example, by past shocks of the spot variance. The theory of storage indeed suggests that past spot variance should affect negatively the conditional covariance.

The model used by Baillie and Myers (1991) is another multivariate GARCH model, labelled BEKK-GARCH model, defined in Engle and Kroner (1995). The bivariate BEKK-GARCH specification is the following:

$$y_t = \mu_t(\eta) + \epsilon_t \tag{2.7}$$

$$\epsilon_t = H_t^{1/2}(\eta) z_t \tag{2.8}$$

$$z_t \rightarrow N(0, I_2) \tag{2.9}$$

$$H_t(\eta) = C'C + A'\epsilon_{t-1}\epsilon'_{t-1}A + B'H_{t-1}B \tag{2.10}$$

where y_t is the vector of spot and futures return (ΔS_t and ΔF_t), $\mu_t(\eta)$ their conditional mean depending on parameters η such as an ARMA process, z_t is the vector of standardized residuals, H_t the $2X2$ covariance matrix of the spot and futures return, A and B are $2X2$ coefficient matrices and C is an upper triangular $2X2$ matrix.

The advantages of the BEKK model are 1) the guarantee, by construction due to the quadratic forms, that the variance-covariance matrix is positive-definite, 2) the time-varying structure of all its components (variances and covariance of spot and futures returns) and 3) the potential cross-variables influences of the spot and futures components (each depend on a constant, on past variances of spot and futures, on past covariance, on past squared spot shock, on past squared futures shock and on cross product of past spot and futures shocks). We also use the BEKK model as a comparison point.

As mentioned in the introduction, the theory of storage predicts different channels of influence of the inventories on the volatility. Using the spot and futures shocks as proxies of the state of the inventory, following in this sense Gorton, Hayashi, and Rouwenhorst (2007), an appropriate multivariate GARCH model should capture the following features:

1. Spot variance should be more reactive to past shocks than futures variances;

2. Spot and Futures covariance should decrease when the spot variance increases;

3. Positive shocks should have a larger impact on the volatility than negative shocks (the inventory effect).

Since the models presented thus far, naïve, OLS and BEKK models, do not allow to measure and capture this last channel of influence, the inventory effect, and given its relative importance as documented in Chapter 1, we use an extension of the BEKK model proposed by Kroner and Ng (1998) where the variance-covariance matrix not only depends on the magnitude of past squared return innovations and variance-covariance matrix but also on the sign of the past squared return innovations. The Asym-BEKK model is the same as the BEKK model, except the covariance-variance matrix which is specified as follows:

$$H_t(\eta) = C'C + A'\epsilon_{t-1}\epsilon'_{t-1}A + B'H_{t-1}B + D'\zeta_{t-1}\zeta'_{t-1}D \tag{2.11}$$

where D is a $2X2$ matrix and ζ_t, the asymmetric term, is a two-variable vector defined as $\zeta_t = \max(\epsilon_t,0)$. Hence, the elements of D represent the effect of past (positive signed) innovations corresponding to the inventory effect. As mentioned above, the Asym-BEKK model is a multivariate transposition of the univariate GJR-GARCH model, developed by Glosten Jagannathan and Runkle (1993) but where the asymmetric term captures the positive shocks instead of the negative ones. To have a full understanding of the channels of interactions, we detail in Equations 2.12, 2.13 and 2.14 the respective equations of the spot variance, futures variance and spot-futures covariance, with Greek letters representing combinations of the elements of coefficient matrices.

$$\begin{aligned} VAR(\Delta S_t) = h_{11,t} &= \gamma_{11} \\ &+\alpha'_{11}\epsilon^2_{1,t-1} + \alpha''_{11}\epsilon_{1,t-1}\epsilon_{2,t-1} + \alpha'''_{11}\epsilon^2_{2,t-1} \\ &+\beta'_{11}h_{11,t-1} + \beta''_{11}h_{12,t-1} + \beta'''_{11}h_{22,t-1} \\ &+\delta'_{11}\zeta^2_{1,t-1} + \delta''_{11}\zeta_{1,t-1}\zeta_{2,t-1} + \delta'''_{11}\zeta^2_{2,t-1} \end{aligned} \tag{2.12}$$

$$\begin{aligned} COV(\Delta S_t, \Delta F_t) = h_{21,t} = h_{12,t} &= \gamma_{12} \\ &+\alpha'_{12}\epsilon^2_{1,t-1} + \alpha''_{12}\epsilon_{1,t-1}\epsilon_{2,t-1} + \alpha'''_{12}\epsilon^2_{2,t-1} \\ &+\beta'_{12}h_{11,t-1} + \beta''_{12}h_{12,t-1} + \beta'''_{12}h_{22,t-1} \\ &+\delta'_{12}\zeta^2_{1,t-1} + \delta''_{12}\zeta_{1,t-1}\zeta_{2,t-1} + \delta'''_{12}\zeta^2_{2,t-1} \end{aligned} \tag{2.13}$$

$$\begin{aligned} VAR(\Delta F_t) = h_{22,t} &= \gamma_{22} \\ &+\alpha'_{22}\epsilon^2_{1,t-1} + \alpha''_{22}\epsilon_{1,t-1}\epsilon_{2,t-1} + \alpha'''_{22}\epsilon^2_{2,t-1} \\ &+\beta'_{22}h_{11,t-1} + \beta''_{22}h_{12,t-1} + \beta'''_{22}h_{22,t-1} \\ &+\delta'_{22}\zeta^2_{1,t-1} + \delta''_{22}\zeta_{1,t-1}\zeta_{2,t-1} + \delta'''_{22}\zeta^2_{2,t-1} \end{aligned} \tag{2.14}$$

In Table 2.1 we provide the correspondence between the Greek letters in Equations 2.12, 2.13 and 2.14 and the components of matrices C, B, A and

D in Equation 2.11. It can be seen that instead of directly interpreting the estimated coefficients, we are referring to their nonlinear combination as the effect of the past variables on the conditional variances and covariance. While the ML estimate of a non-linear function of parameters is calculated as the function of the ML estimate of the parameters, the standard error of the function is obtained by using *"Delta method"*. This method uses the Jacobian matrix of the nonlinear function, evaluated at the estimated parameters.

Table 2.1: "Greek letters" - nonlinear combinations of the estimated coefficients: Asym-BEKK specification

	$h_{11,t}$		$h_{12,t} = h_{21,t}$		$h_{22,t}$
$\gamma_{11} =$	c_{11}^2	$\gamma_{12} =$	$c_{11}c_{12}$	$\gamma_{22} =$	$c_{12}^2 + c_{22}^2$
$\beta_{11}^{'} =$	b_{11}^2	$\beta_{12}^{'} =$	$b_{11}b_{12}$	$\beta_{22}^{'} =$	b_{12}^2
$\beta_{11}^{''} =$	$2b_{11}b_{21}$	$\beta_{12}^{''} =$	$b_{11}b_{22} + b_{12}b_{21}$	$\beta_{22}^{''} =$	$2b_{12}b_{22}$
$\beta_{11}^{'''} =$	b_{21}^2	$\beta_{12}^{'''} =$	$b_{21}b_{22}$	$\beta_{22}^{'''} =$	b_{22}^2
$\alpha_{11}^{'} =$	a_{11}^2	$\alpha_{12}^{'} =$	$a_{11}a_{12}$	$\alpha_{22}^{'} =$	a_{12}^2
$\alpha_{11}^{''} =$	$2a_{11}a_{21}$	$\alpha_{12}^{''} =$	$a_{11}a_{22} + a_{12}a_{21}$	$\alpha_{22}^{''} =$	$2a_{12}a_{22}$
$\alpha_{11}^{'''} =$	a_{21}^2	$\alpha_{12}^{'''} =$	$a_{21}a_{22}$	$\alpha_{22}^{'''} =$	a_{22}^2
$\delta_{11}^{'} =$	d_{11}^2	$\delta_{12}^{'} =$	$d_{11}d_{12}$	$\delta_{22}^{'} =$	d_{12}^2
$\delta_{11}^{''} =$	$2d_{11}d_{21}$	$\delta_{12}^{''} =$	$d_{11}d_{22} + d_{12}d_{21}$	$\delta_{22}^{''} =$	$2d_{12}d_{22}$
$\delta_{11}^{'''} =$	d_{21}^2	$\delta_{12}^{'''} =$	$d_{21}d_{22}$	$\delta_{22}^{'''} =$	d_{22}^2

Notes. $h_{11,t}$, $h_{12,t}$ and $h_{22,t}$ refer to the spot, covariance and futures Equations 2.12, 2.13 and 2.14.

Finally, the log-likelihood function, when the Gaussian distribution is assumed, has the following form:

$$L(\eta) = -\frac{TN}{2}(\ln(2\pi)) - 1/2\sum_{i=1}^{T}(\ln|H_t| + \epsilon_t^{'}H_t^{-1}\epsilon_t^{'}) \tag{2.15}$$

where T is the number of observations and N is the number of variables (series) in the system ($N = 2$ in this paper).

2.4 Data

To illustrate the potential importance of the asymmetries for commodity hedging, we use daily data of the S&P GSCI Index (formerly the Goldman Sachs Commodity Index). This index represents a benchmark measure of performance for the commodity market over time. The index is traded at Chicago Mercantile Exchange. The weight of energy is by far higher than other constituents and it amounts to slightly more that 68%. Other components, as of May 2010, are: Industrial Metals (8.49%), Precious Metals (3.74%), Agriculture (13.7%), and Livestock (5.24%), overall the index is composed of 24 commodities. Because of the broad diversification the index is also interpreted as a measure of general price movements and inflation in the world economy. In this paper we use GSCI Spot Month and GSCI Three Month Future indices, which span the data from January 17, 1995 to April 15, 2010 resulting in 3978 observations. The Data were collected from Datastream database under respective codes CGSYSPT and GS3MSPT.

Table 2.2: Summary statistics

	LEVEL		RETURN	
	Spot	Futures	Spot	Futures
Mean	304.13	174.31	0.028438%	0.028970%
Standard Deviation	152.59	89.215	0.014473	0.012481
Minimum	128.89	76.250	-0.091451	-0.078958
Maximum	890.29	512.61	0.072147	0.070885
Normality test	2296.8**	2329.0**	725.58**	1089.8**
UR test	-2.189	-2.078	-30.27**	-29.82**
$Q^2(20)$ test	77232**	77335**	2575.6**	3698.0**
LM test	87552**	99196**	42.85**	57.38**
Observations	3977	3977	3977	3977

Note: "Level" refers to the value of the index. "Return" refers to the difference of the logarithm of the series in level. The null of the normality test is normality of the series. The null of the unit root (UR) test is the presence of unit root in the series. The null of the Box-Pierce test ($Q^2(20)$) is the absence of autocorrelation in the squared series. The null of the Lagrange Multiplier (LM) test is the absence or ARCH effect in the series. ** stands for rejection at 1%. UR tests on level series based on logarithm of the series do not change the conclusions (not reported here).

The evolution of the two indices over the sample period is shown in Figure 2.1. When tested for unit root, using Augmented Dickey-Fuller (ADF) test, these series are found to be non-stationary. We reiterate the unit root tests on the log-differences of the series and, this time, reject the null. By consequence, the spot and futures series are integrated of order one and taking log-differences make them stationnary. The graph of these daily log-return series is given in Figure 2.2 where both spot and futures follow such a close dynamics that we cannot distinguish one from the other, and in Figure 2.3 where the sample is restricted to the last available year. We can see in this graph that spot return shocks are very often stronger than those of futures returns.

Summary statistics are presented in Table 2.2. As predicted by the theory of storage and documented by Ng and Pirrong (1994), the standard deviation of the spot return is larger than for the futures return. When inventories are large shocks affect both equally spot and futures volatility, but when stocks are low, spot volatility is more affected by a potential stock out than futures

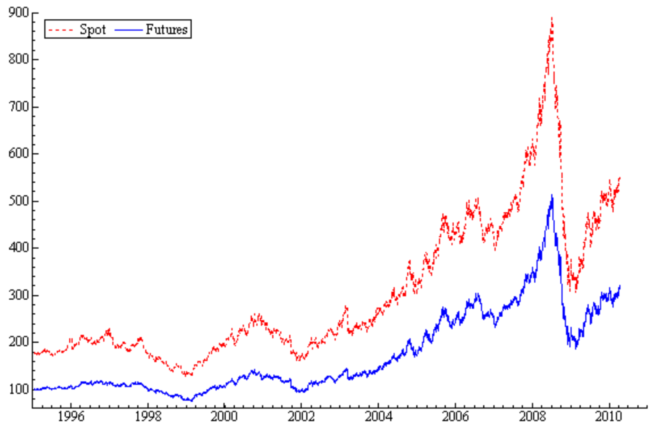


Figure 2.1: Goldman Sachs Commodity Price Index

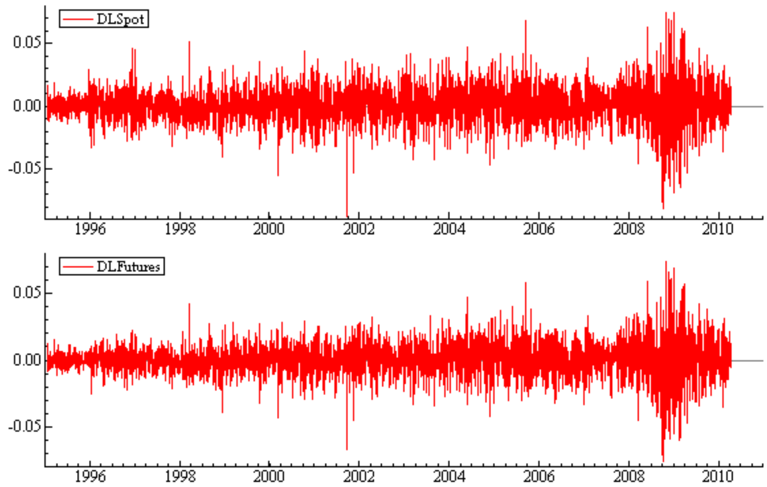


Figure 2.2: Spot and Futures Log-Return Series

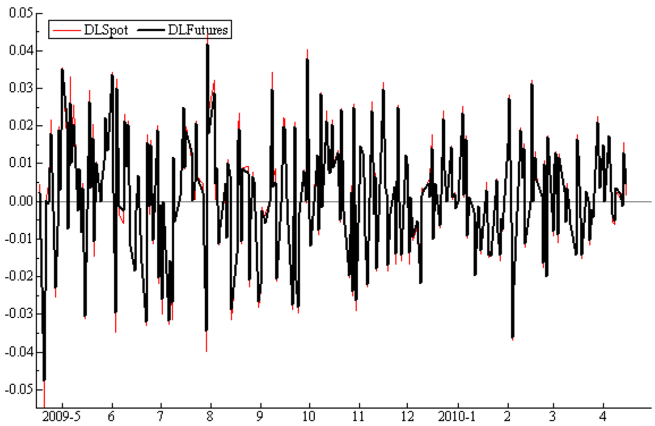


Figure 2.3: Spot and Futures Log-Return Series

for which production has (more) time to adjust. Spot return also has more extreme values (min and max) than futures, following the same logic.

Box-Pierce test on squared series and Largange Multplier test of Engle (1982), reported in Table 2.2, also indicate the presence of heteroskedasticity in the series, which together suggest that using GARCH models would be relevant.

2.5 Results

Our objective is to compare different models in terms of hedging performance and test if adding the asymmetric term to the model delivers better results for hedging a (potential) position in the GSCI index. To do so, we estimate time-varying hedge ratios using two different specifications for the conditional variance-covariance matrix, which are then compared with the constant hedge ratio, obtained from the OLS model and the *naïve* hedge, i.e. the hedge ratio being equal to 1. The *naïve* hedge implies opening an opposite position in futures contracts of the same amount as spot contracts are held at the same time (one-to-one relationship). Below we review the results obtained from estimation of the above models.

The most simple estimate of the OHR is obtained from the OLS model, where it represents simply an estimated slope coefficient, which is constant over time. Table 2.3 shows that the OHR estimated in this way equals to 1.129, meaning that an investor interested in reducing the variance of his/her portfolio would have to sell (buy) 1.129 futures contracts for every spot contract in which (s)he has a long (short) position.

Table 2.3: OLS regression for dependent variable ΔS_t

Exog. Var.	Coef.	Std. Err.	t value	Prob.$>t$
ΔF_t	1.129678	0.004151	272.18	0
Constant	-1.8E-05	2.25E-05	-0.82	0.414

The first panel of Table 2.4 shows the results for the conditional mean Equations, which are specified as AR(1) processes for both spot and futures series. Following Baillie and Myers (1991), we estimate autoregressive models and do not work in a cointegration framework. We indeed found that error correction terms were not significant in the mean equations. [2] P-values are presented in parenthesis and, as can be seen in both cases, the lag coefficient is statistically different from zero, while the constant is insignificant. Instead of this short-term dependence of the spot and futures returns, a random walk was expected according to the efficient market hypothesis. The significance of the lag return is however not an uncommon finding. Lien and Yang (2008), for example, also find significant lags in the mean equations of their commodities. Stock returns also often exhibit short-term dependence, for which potential sources are transaction costs or behaviorial patterns (as discussed in the review of Malkiel (2003) or the chapter on predictability of asset returns in Campbell and MacKinlay (1996)).

Panel two of Table 2.4 displays the coefficient estimates for the conditional variance-covariance matrix. The elements of A and B matrices represent the ARCH and GARCH effects respectively and C is an upper-triangular matrix representing the constant term. P-values in parenthesis suggest high statistical significance of both ARCH and GARCH effects.

The results for the Asym-BEKK model, given in Table 2.5, suggest the presence of asymmetries in the conditional covariance structure. Particularly, all elements of matrix D are statistically significant even at 1% level.

[2] Results available on request from the authors.

Table 2.4: BEKK-GARCH Model Estimates

Conditional Mean Equations (P-values in parenthesis):

$$y_{1,t} = \underset{(0.61)}{0.000040} + \underset{(0.00)}{0.102380} y_{1,t-1} + \epsilon_{1,t}$$

$$y_{2,t} = \underset{(0.23)}{0.000082} + \underset{(0.00)}{0.108843} y_{2,t-1} + \epsilon_{2,t}$$

Conditional Variance-Covariance Matrix (P-values in parenthesis):

$$H_t(\eta) = C'C + A'\epsilon_{t-1}\epsilon'_{t-1}A + B'H_{t-1}B$$

$$C = \begin{bmatrix} 0.000740 & 0.000439 \\ (0.00) & (0.00) \\ 0 & 0.00001 \\ & (0.00) \end{bmatrix} \quad A = \begin{bmatrix} -0.851078 & -0.405689 \\ (0.00) & (0.00) \\ 1.031991 & 0.551631 \\ (0.00) & (0.00) \end{bmatrix}$$

$$B = \begin{bmatrix} 0.813052 & -0.070711 \\ (0.00) & (0.00) \\ 0.191031 & 1.067519 \\ (0.00) & (0.00) \end{bmatrix}$$

Table 2.5: Asym-BEKK Model Estimates

Conditional Mean Equations (P-values in parenthesis):

$$y_{1,t} = \underset{(0.45)}{0.000141} + \underset{(0.00)}{0.113007} y_{1,t-1} + \epsilon_{1,t}$$

$$y_{2,t} = \underset{(0.11)}{0.000262} + \underset{(0.00)}{0.116116} y_{2,t-1} + \epsilon_{2,t}$$

Conditional Variance-Covariance Matrix (P-values in parenthesis):

$$H_t(\eta) = C'C + A'\epsilon_{t-1}\epsilon'_{t-1}A + B'H_{t-1}B + D'\zeta_{t-1}\zeta'_{t-1}D$$

$$C = \begin{bmatrix} 0.000734 & 0.000327 \\ (0.00) & (0.00) \\ 0 & -0.000003 \\ & (0.00) \end{bmatrix} \quad A = \begin{bmatrix} -0.401618 & -0.045789 \\ (0.00) & (0.15) \\ 0.243888 & -0.127907 \\ (0.00) & (0.00) \end{bmatrix}$$

$$B = \begin{bmatrix} 0.921622 & -0.003593 \\ (0.00) & (0.69) \\ 0.066443 & 0.986246 \\ (0.00) & (0.00) \end{bmatrix} \quad D = \begin{bmatrix} -0.348236 & -0.206677 \\ (0.00) & (0.00) \\ 0.479650 & 0.328875 \\ (0.00) & (0.00) \end{bmatrix}$$

However, as mentioned earlier, the BEKK-GARCH specification does not allow a direct interpretation of these elements as the impact of past variables on the contemporaneous volatility. Instead, we are interested in interpreting combinations of these parameters, which are reported in Table 2.6. The standard errors, reported in the columns next to the coefficients, were estimated using the delta method. The correspondence between Tables 2.5 and 2.6 is detailed in Table 2.1.

First, past shocks, ϵ_1^2 and ϵ_2^2 , have a larger impact on the spot series than on the futures one, which confirms the largest reactivity of spot prices. Second, when spot and futures shocks don't have opposite signs (which would signal potential inventory shortage), the spot variance decreases (coefficient of $\epsilon_1\epsilon_2$ negative). Third, all asymmetric coefficients are significant. Fourth, past positive spot and futures shocks have a positive impact on both variances, which illustrates the inventory effect. Finally, when both spot and futures shocks are positive at the same time, the spot and futures variances decrease.

Table 2.6: Estimated nonlinear coefficients and standard errors for Asym-BEKK Model using Delta Method

	Equation $h_{11,t}$		Equation $h_{12,t}$		Equation $h_{22,t}$	
	Coefficient	St. Error	Coefficient	St. Error	Coefficient	St. Error
γ	1.0E-5***	1.3E-5	1.4E-7**	2.1E-7	2.9E-6	1.0E-6
$\alpha^{'}$	0.161***	0.045	0.018	0.015	0.002	0.003
$\alpha^{''}$	-0.196**	0.082	0.040*	0.022	0.012***	0.004
$\alpha^{'''}$	0.059*	0.034	-0.031***	0.005	0.016	0.011
$\beta^{'}$	0.849***	0.030	-0.003	0.008	1.3E-6	6.5E-6
$\beta^{''}$	0.122***	0.031	0.909***	0.011	-0.007	0.018
$\beta^{'''}$	0.004*	0.002	0.066***	0.019	0.972***	0.022
$\delta^{'}$	0.121*	0.064	0.072**	0.033	0.043**	0.022
$\delta^{''}$	-0.334***	0.149	-0.214**	0.086	-0.136**	0.064
$\delta^{'''}$	0.230***	0.088	0.158***	0.059	0.108**	0.050

Note: Triple, double and single asterisk represents significance at 1, 5 and 10 per cent levels respectively.

Figure 2.4 graphs the time-varying OHRs as computed from the symmetric and Asym-BEKK models, as well as OLS and *naïve* hedge ratios. The average OHR for the asymmetric model is slightly lower than that obtained from the symmetric model (1.162967 and 1.16468 respectively) and both are higher than OLS hedge ratio. The two time-varying ratios are mostly above one, implying that more than one futures contract should be traded for each spot index contract[3].

[3] Augmented Dickey-Fuller (ADF) tests suggest that the time-varying ratios are stationary.

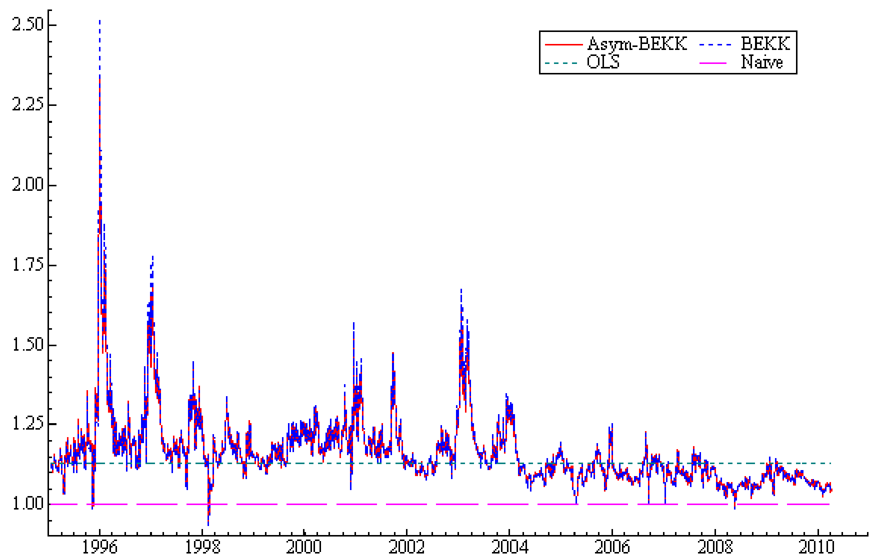


Figure 2.4: In-sample forecasts - Optimal Hedge Ratios

To compare in-sample hedging performance we take the full sample. The four OHRs are used to choose optimal positions in futures contracts over the sample period. Then, the sample variance evaluated for the whole period is used as the measure of hedging performance. The naïve hedge is the worst model for the in-sample case. Using the OLS approach reduces the portfolio variance by about 20% and the time-varying OHR by about 27%. Using the Asym-BEKK provides the best performance (the lowest variance) as reported in Table 2.7.

Next, we compare the out-of-sample performances. The one-day ahead forecasting is performed for 150 days (observations). For each $t + 1$ forecast, the model is reestimated by using the data available up to date t. Based on the variance-covariance matrix estimate, the expected OHR is used to build the portfolio (long in the spot and short in the futures up to the number of contracts determined by the OHR) and compute its return and variance. We start at the 3828th observation and repeat the procedure until the 3977th observation. The generated series are illustrated in Figure 2.5.

The results demonstrate that the asymmetric model yields superior hedging performance relative to all other models. As can be seen from Table 2.8 the portfolio variance is the lowest for the Asym-BEKK, 16% below the

Table 2.7: In-sample forecasts: variance of naïve, OLS, BEKK and Asym-BEKK portfolios

	naïve	OLS	BEKK	Asym-BEKK
	$\beta = 1$	$\beta = 1.1296$	$\beta = h_{12,t}/h_{22,t}$	$\beta = h_{12,t}/h_{22,t}$
Variance	0.0133	0.0107	0.0097	0.0097
% of naïve Variance	100.00%	80.38%	73.38%	73.02%

Notes. For reporting purposes variance are multiplied by 1,000. The row "% of naïve Variance" reports the variances of portfolio returns for the different strategies as a percentage of the naïve variance.

Figure 2.5: Out-of-sample forecasts - Optimal Hedge Ratios

BEKK portfolio variance, and 6.55% below the naïve hedge which performs surprinsingly well compared to the BEKK.

Table 2.8: Out-of-sample forecasts: variance of naïve, OLS, BEKK and Asym-BEKK portfolios

	naïve	OLS	BEKK	Asym-BEKK
	$\beta = 1$	$\beta = 1.1335$	$\beta = h_{12,t}/h_{22,t}$	$\beta = h_{12,t}/h_{22,t}$
Variance	0.0010	0.0017	0.0011	0.0009
% of naïve Variance	100.00%	176.86%	111.37%	93.46%

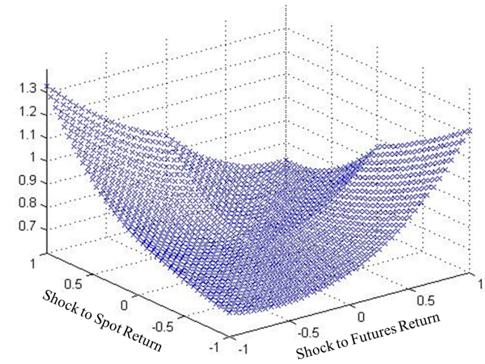
Notes. For reporting purposes variance are multiplied by $1,000$. The row "% of naïve Variance" reports the variances of portfolio returns for the different strategies as a percentage of the naïve variance.

The results demonstrate that the asymmetric model yields superior hedging performance relative to all other models. As can be seen from Table 2.7 the portfolio variance is the the lowest for the Asym-BEKK, 16% below the BEKK-GARCH portfolio variance, and 6.55% below the *naïve* hedge which performs surprinsingly well compared to the BEKK-GARCH. Finally, the mean portfolio return is also highest for the portfolio hedged using the asymmetric multivariate model.

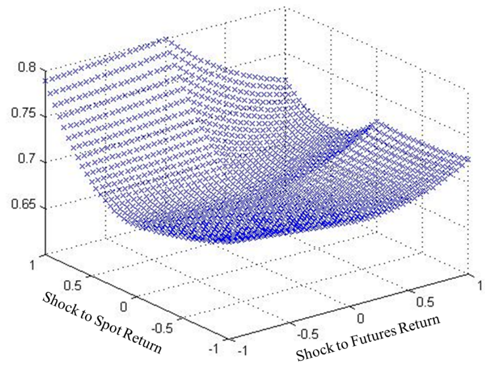
In their paper, Brooks, Henry, and Persand (2002) suggest linking the concept of the OHR with the notion of the *News Impact Surface* developed by Kroner and Ng (1998). According to Kroner and Ng (1998) changes in prices can be viewed as additional information available to the market. Therefore, such innovations in prices, $S_t - S_{t-1} = \epsilon_{S,t}$ and $F_t - F_{t-1} = \epsilon_{F,t}$, represents a measure of news arriving to the market between periods t and $t-1$. Consequently, the relationship between innovations in returns and the conditional variance-covariance structure is defined as the news impact surface. The OHR β^*_{t-1}, as defined in Equation 2.6, can be linked to news, by using news impact surfaces of futures variance and spot-futures covariance. The resulting news impact surface may be interpreted as the response of the OHR to arrival of the news.

Figure 2.6 plots the news impact surfaces for the conditional spot and futures variances, and the conditional spot-futures covariance, which are obtained from the Asym-BEKK model. Each surface is evaluated in the region $\epsilon_{j,t}$ =[-1, 1], with $j = Spot, Futures$.

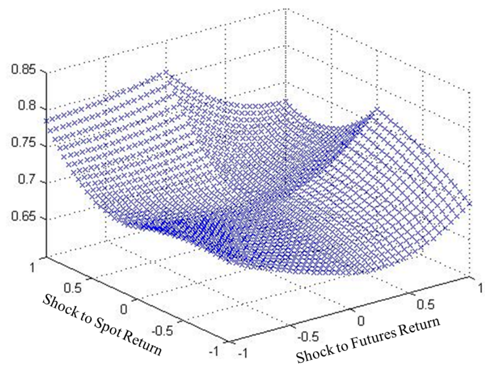
The asymmetric response of the variances and covariance is clearly seen from



(a) Spot Variance



(b) Futures Variance



(c) Spot-Futures Covariance

Figure 2.6: News Impact Surfaces

these figures. Particularly, the conditional spot variance reaches slightly higher levels in magnitude when both shocks are positive compared to the case where both are negative.

Moreover, we see that opposite (in sign) values of the shocks largely increase the conditional variance of the spot (left and right side of upper graph in Figure 2.6 are higher). This is not surprising since it reveals short-term shortages. On the contrary, periods where spot and futures dynamics coincide correspond to periods of large inventories and of relative quietness on the commodity markets.

In addition, large shocks in the spot returns contribute more to the conditional variance than large shocks in the futures return. The extreme left side of the Spot variance NIS in Figure 2.6 is higher than the extreme right side. This is also an implication of the theory of storage by which spot shocks better proxy the state of the inventories. A similar pattern, though weaker, is found for the news impact surface of the futures variance, i.e. more dramatic response of the variance is found for the region where spot return innovations are positive.

Lien and Yang (2008) tested on 10 commodities (all others than ours) whether taking into account asymmetric inventory effect improves hedging strategies. They used the basis as proxy of the inventory effect and modelled the dynamic covariance through a DCC model. They found that inventories affect negatively the volatility for all but one commodity (cotton). We also find this negative effect for all but one commodity (nickel). Regarding hedging performance, they find that their asymmetric approach reduces the variance by up to 3% for in-sample performance. We find with our approach similar improvements with variance reduction up to 4%. However, our results cannot be directly compared to those of Lien and Yang (2008). Firstly, the samples and commodities are different. Secondly, we use as proxy positive return shocks, while they use the basis. Finally, we use a more flexible specification for the dynamic covariance, where the covariance is not only affected by its own past, but also by asymmetric terms. This makes our approach less restictive and potentially more performant.

2.6 Extensions

We propose two extensions. First, we check if taking into account the inventory effect through an Asym-BEKK model also improves the OHR performance for some other commodities. We then consider an alternative optimality framework not restricted to the variance reduction. We more precisely relax the martingale assumption and maximize the (quadratic) expected utility by taking into account the expected return.

2.6.1 Other commodities

To check the relevance of using an Asym-BEKK model to estimate a dynamic OHR, we reiterate our in-sample and out-of-sample forecasts for 3 additional commodities: lead, nickel and aluminium[4]. These metal price series are available as daily observations on Datastream under labels LEDCASH/LED3MTH for lead, LNICASH/LNI3MTH for nickel and LAHCASH/LAH3MTH for aluminium. The suffix 'cash' and '3mth' refer to spot and futures, respectively. The advantage of using these series, provided by the London Metal Exchange, is their availability in a continuous form (no need to compile successive futures contracts and correct for rollover discrepancies). Another advantage is their availability on the same time horizons as our benchmark S&P GSCI Index (spot and 3-month horizons).

We find in Table 2.9 that the Asym-BEKK models also dominate the BEKK-GARCH model in terms of hedging performance for two other commodities (lead and aluminium). As regards lead, the variance of the hedged portfolio is reduced by 4.6% and 17.8% for in-sample and out-of-sample forecasts, respectively. The Asym-BEKK also improves hedging for aluminium in similar proportions. However, using an Asym-BEKK does not improve the hedging performance for nickel. On the contrary, it marginally increases the variance. These results confirm the relevance of checking for potential inventory effect in view to upgrade the commodity risk management.

Table 2.9: In- and Out-of-sample forecasts: Portfolio variance reduction obtained by using an Asym-BEKK instead of a BEKK model

Commodity	in-sample	out-of-sample
Lead	-4.56%	-17.77%
Nickel	+0.15%	+2.45%
Aluminium	-4.52%	-28.25%

Note: The OHR used in this forecast exercice is β^*_{t-1} (see Equation 2.5). Variance reduction is computed by dividing the variance of the spot-futures portfolio resulting from an Asym-BEKK modelling strategy on the variance of the portfolio resulting from a BEKK modelling strategy.

[4] The computing time for out-of-sample forecasts is considerable. Each series and each model requires 150 estimations.

2.6.2 Mean-variance framework

The OHR depends on the particular objective function to be optimized. Minimizing the variance is the most widely used approach, but there exist alternatives. Chen, Lee, and Shrestha (2003) and Lien and Tse (2002) present a review of different theoretical approaches to OHR and list, as alternatives to the variance minimization, the mean-variance, the expected utility, the mean extended-Gini coefficient as well as the semivariance approaches. These alternatives all consider a priori that minimizing the variance is too restrictive and that the risk aversion or the expected return should be taken into account. Actually, most of these approaches are equivalent to variance minimization under (reasonable) assumptions.

A key assumption, on which we relied in Section 2.2 to derive the OHR from Equation 2.4, was that the futures return follows a martingale, i.e. $E_{t-1}(\Delta F_t) = 0$. Under this assumption, the mean-variance framework (whose objective function to maximize is $E(R_p) - \theta Var(R_p)$, where θ is the risk aversion and R_p the portfolio return) is equivalent to the variance minimization. Even optimizing the expected-utility, where the objective function to maximize is $E(U(Wealth))$, turns out to be equivalent to the variance minimization if the utility function is quadratic (as in Kroner and Sultan (1991)).

Assuming that futures return follows a martingale has some empirical grounds. Indeed, the constants in the mean equations reported in Tables 2.4 and 2.5 are not significant, this tends to give support to this assumption. Notwithstanding this empirical indication, we now relax it to check the robustness of our results. A first reason is that the return on the portfolio hedged according to the Asym-BEKK model does not dominate the return of the portfolio hedged following the BEKK-GARCH modelling. In some cases, the portfolio return is even lower. Improving the risk management performance in this case costs a lower return. The second reason why we check our findings in a mean-variance framework is to implictly check if the martingale assumption is reasonable. Radically different results, compared to Section 2.5 would seriously challenge the martingale assumption.

For convenience, we rewrite here the quadratic utility function to maximize (cf Equation 2.4):

$$\begin{aligned} max \quad & E_{t-1}(\Delta S_t) - \beta_{t-1} E_{t-1}(\Delta F_t) \\ & -\theta(VAR(\Delta S_t) + \beta_{t-1}^2 VAR(\Delta F_t) - 2\beta_{t-1} COV(\Delta S_t, \Delta F_t)) \end{aligned} \tag{2.16}$$

If we relax the martingale assumption for the futures return and solve this mean-variance optimization problem, we now find that the OHR, β_{t-1}^{**} has an additional term:

$$\beta_{t-1}^{**} = \frac{COV(\Delta S_t, \Delta F_t)}{VAR(\Delta F_t)} - \frac{E_{t-1}(\Delta F_t)}{2\theta VAR(\Delta F_t)} \tag{2.17}$$

Following Lien and Yang (2008) and the empirical estimation of Chou (1988), we assume that θ, the degree of risk aversion, is equal to 4 and compare our strategies performances. We report our results in terms of utility in Table 2.10 and find that the Asym-BEKK model again dominates the BEKK modelling of the variance-covariance matrix even when a mean component is considered in the objective function. This is also the case now for nickel, while we found in the variance minimization approach that the Asym-BEKK did not improve the risk performance for this commodity. The respective estimates of the BEKK-GARCH and Asym-BEKK models for the aluminium, nickel and lead series are reported in Tables 2.11, 2.12, 2.13, 2.14, 2.15 and 2.16 in the appendix of this chapter.

Table 2.10: Mean-variance framework: In/Out-of-sample forecasts - Portfolio utility increase obtained by using an Asym-BEKK instead of a BEKK model

Commodity	In-sample	Out-of-sample
GSCI Index	0.0%	+0.1%
Lead	+6.1%	+12.4%
Nickel	+6.9%	+20.7%
Aluminium	+0.4%	+20.4%

Note: The OHR used in this forecast exercice is β^{**}_{t-1} (see Equation 2.17). Utility is defined as $E(R_p) - \theta Var(R_p)$. Utility increase is computed by comparing the utility of the spot-futures portfolio resulting from an Asym-BEKK modelling strategy on the utility of the portfolio resulting from a BEKK modelling strategy.

2.7 Conclusions

Commodities are often found to exhibit an inverse leverage effect. In other words the response of the conditional variance to past negative shocks is weaker than the response to positive shocks of the same magnitude, i.e. opposite to the "leverage effect" documented for equity data in the empirical finance literature. This paper aims to show that addressing optimal hedging problem in a multivariate setup that allows asymmetric response of the conditional variance-covariance matrix delivers a superior measure of OHR and reduces the risk that investors face when hedging their positions in spot contracts using futures contracts. To account for such asymmetries we estimate a flexible, asymmetric multivariate GARCH model in which the covariance structure is specified in such a way that past positive and negative shocks have different effect on the conditional variances and covariance.

Our results confirm that such an asymmetric model yields improvements in both in-sample and out-of-sample forecasting and reduces the variance of realized returns on a portfolio consisting of spot and futures contracts on a broad commodity index (GSCI). Robustness checks operated on 3 other commodities and in a mean-variance optimization framework provide results confirming the relevance to consider the inventory effect. These findings imply that using asymmetric multivariate models would be more effective in hedging compared with those that assume symmetric response to past innovations. In addition to hedging optimization, the results have important implications for Value-at-Risk applications, as using the asymmetric model would improve assessment of this widespread risk measure.

Appendix: BEKK-GARCH and Asym-BEKK Estimates for the aluminium, nickel and lead series

Table 2.11: BEKK-GARCH Model Estimates for Aluminium

Conditional Mean Equations (P values in parenthesis):

$$y_{1,t} = \underset{(0.97)}{0.000006} + \underset{(0.00)}{0.120651} y_{1,t-1} + \epsilon_{1,t}$$

$$y_{2,t} = \underset{(0.88)}{0.000025} + \underset{(0.00)}{0.117683} y_{2,t-1} + \epsilon_{2,t}$$

Conditional Variance-Covariance Matrix (P values in parenthesis):

$$H_t(\eta) = C'C + A'\epsilon_{t-1}\epsilon'_{t-1}A + B'H_{t-1}B$$

$$C = \begin{bmatrix} 0.001244 & 0.000867 \\ (0.00) & (0.00) \\ 0 & 0.000103 \\ & (0.00) \end{bmatrix} \quad A = \begin{bmatrix} -0.476768 & -0.046696 \\ (0.00) & (0.43) \\ 0.634730 & 0.185678 \\ (0.00) & (0.00) \end{bmatrix}$$

$$B = \begin{bmatrix} 0.915819 & 0.000768 \\ (0.00) & (0.96) \\ 0.071633 & 0.987102 \\ (0.01) & (0.00) \end{bmatrix}$$

Table 2.12: Asym-BEKK Model Estimates for Aluminium

Conditional Mean Equations (P values in parenthesis):

$$y_{1,t} = \underset{(0.54)}{0.000106} + \underset{(0.00)}{0.124773} y_{1,t-1} + \epsilon_{1,t}$$

$$y_{2,t} = \underset{(0.50)}{0.000112} + \underset{(0.00)}{0.122044} y_{2,t-1} + \epsilon_{2,t}$$

Conditional Variance-Covariance Matrix (P values in parenthesis):

$$H_t(\eta) = C'C + A'\epsilon_{t-1}\epsilon'_{t-1}A + B'H_{t-1}B + D'\zeta_{t-1}\zeta'_{t-1}D$$

$$C = \begin{bmatrix} 0.001526 & 0.001302 \\ (0.00) & (0.00) \\ 0 & 0.000266 \\ & (0.40) \end{bmatrix} \quad A = \begin{bmatrix} -0.400325 & -0.001138 \\ (0.00) & (0.99) \\ 0.542842 & 0.122241 \\ (0.00) & (0.12) \end{bmatrix}$$

$$B = \begin{bmatrix} 0.933059 & 0.013103 \\ (0.00) & (0.64) \\ 0.040077 & 0.960979 \\ (0.17) & (0.00) \end{bmatrix} \quad D = \begin{bmatrix} 0.074625 & -0.108590 \\ (0.48) & (0.18) \\ 0.157953 & 0.333796 \\ (0.23) & (0.00) \end{bmatrix}$$

Table 2.13: BEKK-GARCH Model Estimates for Lead

Conditional Mean Equations (P values in parenthesis):

$$y_{1,t} = \underset{(0.57)}{0.000126} + \underset{(0.00)}{0.166647} y_{1,t-1} + \epsilon_{1,t}$$

$$y_{2,t} = \underset{(0.42)}{0.000165} + \underset{(0.00)}{0.167800} y_{2,t-1} + \epsilon_{2,t}$$

Conditional Variance-Covariance Matrix (P values in parenthesis):

$$H_t(\eta) = C'C + A'\epsilon_{t-1}\epsilon'_{t-1}A + B'H_{t-1}B$$

$$C = \begin{bmatrix} 0.002910 & 0.001273 \\ (0.00) & (0.00) \\ 0 & 0.000000 \\ & (0.00) \end{bmatrix} \quad A = \begin{bmatrix} -0.694243 & 0.178799 \\ (0.00) & (0.00) \\ 0.848699 & 0.313452 \\ (0.00) & (0.00) \end{bmatrix}$$

$$B = \begin{bmatrix} 0.809674 & -0.032833 \\ (0.00) & (0.13) \\ 0.179680 & 1.022947 \\ (0.00) & (0.00) \end{bmatrix}$$

Table 2.14: Asym-BEKK Model Estimates for Lead

Conditional Mean Equations (P values in parenthesis):

$$y_{1,t} = \underset{(0.00)}{0.000480} + \underset{(0.00)}{0.165828} y_{1,t-1} + \epsilon_{1,t}$$

$$y_{2,t} = \underset{(0.00)}{0.000410} + \underset{(0.00)}{0.166435} y_{2,t-1} + \epsilon_{2,t}$$

Conditional Variance-Covariance Matrix (P values in parenthesis):

$$H_t(\eta) = C'C + A'\epsilon_{t-1}\epsilon'_{t-1}A + B'H_{t-1}B + D'\zeta_{t-1}\zeta'_{t-1}D$$

$$C = \begin{bmatrix} 0.003280 & 0.001799 \\ (0.00) & (0.00) \\ 0 & 0.000001 \\ & (0.14) \end{bmatrix} \quad A = \begin{bmatrix} -0.514071 & -0.143532 \\ (0.00) & (0.01) \\ 0.689060 & 0.285643 \\ (0.00) & (0.00) \end{bmatrix}$$

$$B = \begin{bmatrix} 0.829764 & -0.035877 \\ (0.00) & (0.21) \\ 0.140808 & 1.013492 \\ (0.02) & (0.00) \end{bmatrix} \quad D = \begin{bmatrix} 0.344528 & -0.074043 \\ (0.00) & (0.26) \\ -0.116299 & 0.275064 \\ (0.33) & (0.00) \end{bmatrix}$$

Table 2.15: BEKK-GARCH Model Estimates for Nickel

Conditional Mean Equations (P values in parenthesis):

$$y_{1,t} = 0.001056 + 0.183749 y_{1,t-1} + \epsilon_{1,t}$$
(0.02) (0.00)
$$y_{2,t} = 0.000879 + 0.180413 y_{2,t-1} + \epsilon_{2,t}$$
(0.04) (0.00)

Conditional Variance-Covariance Matrix (P values in parenthesis):

$$H_t(\eta) = C'C + A'\epsilon_{t-1}\epsilon'_{t-1}A + B'H_{t-1}B$$

$$C = \begin{bmatrix} 0.001671 & 0.001218 \\ (0.00) & (0.00) \\ 0 & 0.000000 \\ & (0.21) \end{bmatrix} A = \begin{bmatrix} 0.248242 & -0.348764 \\ (0.02) & (0.02) \\ -0.109046 & 0.424299 \\ (0.25) & (0.00) \end{bmatrix}$$

$$B = \begin{bmatrix} 0.961257 & 0.131492 \\ (0.00) & (0.07) \\ 0.027146 & 0.860329 \\ (0.45) & (0.00) \end{bmatrix}$$

Table 2.16: Asym-BEKK Model Estimates for Nickel

Conditional Mean Equations (P values in parenthesis):

$$y_{1,t} = 0.001056 + 0.195484 y_{1,t-1} + \epsilon_{1,t}$$
(0.00) (0.00)
$$y_{2,t} = 0.000886 + 0.191546 y_{2,t-1} + \epsilon_{2,t}$$
(0.00) (0.00)

Conditional Variance-Covariance Matrix (P values in parenthesis):

$$H_t(\eta) = C'C + A'\epsilon_{t-1}\epsilon'_{t-1}A + B'H_{t-1}B + D'\zeta_{t-1}\zeta'_{t-1}D$$

$$C = \begin{bmatrix} 0.001365 & 0.001761 \\ (0.00) & (0.00) \\ 0 & -0.000004 \\ & (0.00) \end{bmatrix} A = \begin{bmatrix} 0.490901 & -0.083505 \\ (0.00) & (0.59) \\ -0.352301 & 0.158567 \\ (0.00) & (0.22) \end{bmatrix}$$

$$B = \begin{bmatrix} -0.136157 & 0.677348 \\ (0.75) & (0.06) \\ 1.148693 & 0.293917 \\ (0.01) & (0.43) \end{bmatrix} D = \begin{bmatrix} -0.181427 & 0.080674 \\ (0.27) & (0.63) \\ 0.333284 & 0.065307 \\ (0.04) & (0.70) \end{bmatrix}$$

Chapter 3

Real Exchanges Rates in Commodity Producing Countries: A Reappraisal[1]

Abstract

Commodity price booms, as those recorded in the last decade, may have a significant economic impact in small, commodity exporting, developing countries. Whether the impact on output is positive or negative is still unclear. It depends on various factors, notably on the impact that commodity prices can have on the real exchange rate of the commodity exporting countries. Two recent papers show that the real exchange rate appreciates when commodity prices increase. Our analysis produces new estimates of this relationship by focusing on a large sample of developing countries which are specialized in the export of one leading commodity. By using non-stationary panel techniques robust to cross-sectional-dependence, we find that the price of the dominant commodity has a significant long-run impact on the real exchange rate when the exports of the leading commodity have a share of at least 20 percent in the country's total exports of merchandises. Our results also show that the larger this share, the larger the size of the impact.

Keywords: Commodity producers * Commodity prices * Natural resource curse * Non-stationary panel * Real exchange rates

JEL Classification: C32, C33, E31, F32, 011.

[1]Joint work with Vincent Bodart and Bertrand Candelon. Forthcoming in the Journal of International Money and Finance.

3.1 Introduction

Many developing countries, gifted in natural resources, heavily depend on international commodity prices. Whether natural resources are a luck or a curse for these countries is still an open debate. Historical experience shows indeed that some countries have admirably well taken profits of their natural endowments, while others suffered countless economic difficulties. The "natural resource curse" literature, surveyed recently by Frankel (2010b), identifies several channels through which commodity prices can affect the economic performance of commodity producing countries. One important channel is the real exchange rate. The theoretical literature indeed shows that commodity price booms may cause an appreciation of the real exchange rate of commodity exporting countries. This appreciation in turn generates so-called "Dutch disease" effects, by altering the competitiveness of the non-commodity exportable sectors. In this paper, we explore this "real exchange rate" channel and we provide new empirical estimates of the relationship between commodity prices and real exchange rates for a large group of developing countries.

Recent estimates of the relationship that may exist between real exchange rates and commodity prices are provided by Chen and Rogoff (2003) and by Cashin, Cespedes, and Sahay (2004). The former provides results for three developed countries (Canada, Australia and New-Zealand) and the latter for about sixty developing countries. Both use time-series cointegration techniques and find that commodity price increases have a significant positive long-run impact on the real exchange rate of many countries. Our analysis is similar in spirit to those two recent studies. One main difference with them comes from the fact that we focus exclusively on developing countries whose exports are highly concentrated on one particular commodity. In their studies, Chen and Rogoff (2003) and Cashin, Cespedes, and Sahay (2004) both look at the long-run dependence that may exist between a country's real exchange rate and a specific country commodity price index constructed as a weighted average of the price of the different commodities that are exported by the country. In our study however, what we are interested in is to assess whether the price of the leading commodity exported by a country is an economically and statistically significant determinant of the long-run variations in their real exchange rate. The price variable in our cointegration analysis is therefore a single price rather than a constructed price index. We believe that this approach is particularly relevant given the numerous countries that are heavily specialized in the production and export of one natural commodity. This is clearly the case of most oil producing countries. This is also true for many countries exporting agricultural or mineral commodities, as for instance Central African Republic (cotton), Chile (copper), Colombia (coffee), Ivory Coast (cocoa), Ethiopia (coffee), Malawi (tobacco), Mauritius (sugar) or Niger (uranium). More formally, in a very recent paper, Easterly and Reshef (2010) show from a sample of 37 African countries that the cross-country mean share of the top export is about 48 percent of total export receipts (for the period 1994-2008), while the average share of the second top export is only 14 percent of total export receipts. Similar conclusions are drawn from the data reported by Cashin, Cespedes, and Sahay (2004), which covers 58 developing countries[2]. It appears in particular that for 27 countries, the leading commodity accounts for more than 50 percent of total commodity exports, the cross-country mean share being about 75 percent.

For small countries specialized in the export of one main commodity, fluctuations in the (world) price of their leading commodity export can be an important source of macroeconomic instability. For that reason, the price of that commodity can play a key role in the macroeconomic management of these countries. It can for instance influence the design of exchange rate and monetary policies if policymakers are willing to reduce the undesirable "Dutch disease" effects caused by commodity price booms. On this matter, Frankel (2010a) reports that the 2003-2008 boom in world commodity markets put upward pressure on the currency of several Latin American commodity exporting countries and these countries intervened heavily on the foreign exchange market to dampen the currency appreciation. Colombia was one of these countries, as illustrated by Vargas (2005). Edwards (1986) also shows that, in the past, Colombian authorities have also designed schemes to reduce the impact of changes in the world price of coffee on the real exchange rate and the rest of the economy[3]. Regarding other macroeconomic policies, Chile offers an interesting example of how the price of one leading commodity can influence the design of fiscal policy. Frankel and Saiki (2002) reports, under the Chilean rules, that the extent to which the government budget deficit is allowed to deviate from a predetermined target is partly dependent on the price of copper.

Our analysis is obviously not the first one that is looking at the role that one leading commodity can play in the determination of the real exchange of commodity dependent countries. Several studies have for instance shown that the world price of oil is the major determinant of the real exchange rate of oil dependent countries[4]. Regarding non-oil commodities, there is however not plenty of evidence. One study is Edwards (1986), who find that the world price of coffee is a significant explanatory variable of the real exchange rate of Colombia, a world leading exporter of coffee. Other studies are Aron,

[2] From the data reported by Cashin, Cespedes, and Sahay (2004) in their Table 1 (pp.246-247), it appears that the top commodity export accounts on average (across countries and for the period 1991-1999) for 52 percent of all commodity exports. The average share of the second and third leading commodity is respectively about 18 percent and 9 percent.

[3] In particular, Edwards (1986) reports that in 1967 Colombia adopted a crawling peg to reduce the dependence of the real exchange rate on coffee prices fluctuations, and the decision on the rate at which the peso should be devalued was highly influenced by increases in coffee prices

[4] See for instance Habib and Kalamova (2007).

Elbadawi, and Kahn (1998) and Macdonald and Ricci (2004), which provide evidence of a significant positive dependence between the real exchange rate of South Africa, a major world producer of gold, and the price of gold[5]. Most of the existing studies are concerned with one particular country. In contrast, our study uses a large set of developing countries as the one used by Cashin, Cespedes, and Sahay (2004). To our knowledge, it is therefore the first study to examine for a large group of countries the role played by the price of one leading commodity in the determination of the real exchange rate of countries with one dominant exportable commodity. As we use a large number of countries, we are also able to investigate whether the relationship between the real exchange rate of countries specialized in the export of one particular commodity and the price of that commodity is changing with the degree of commodity specialization.

Panel cointegration techniques offer a very convenient setup to tackle the issues mentioned above. We use non-stationary panel cointegration methods of second generation, which are robust to cross-sectional dependence. Anticipating on our results, we find that the price of the leading commodity is a long-run determinant of the real exchange rate of countries where the main commodity accounts for at least 20 percent of the total merchandise exports of the country. We also find that the higher the specialization of the country, the larger the elasticity.

As a byproduct, we also show that the estimate of the cointegration relationship between real exchange rates and commodity prices is markedly different whether the econometric technique that is used is robust or not to cross-section dependence. Real exchange rates are by construction interdependent and commodity prices share some co-movements, as confirmed by Pesaran (2004) tests. By comparing estimates obtained from robust and non-robust techniques, from Bai, Kao, and Ng (2009) on the one hand and fully modified OLS and dynamic OLS on the second hand, we find that the elasticity of real exchange rates to commodity prices is strongly overestimated when methods not robust to cross-sectional dependence are used, as it is the case in Coudert, Couharde, and Mignon (2008).

The rest of the paper is organized as follows. Section 3.2 presents briefly the empirical literature devoted to the relation between commodity prices and real exchange rates. Section 3.3 presents the model. Section 3.4 describes the data. Section 3.5 presents the methods. Section 3.6 discusses the empirical results. Some robustness checks are performed in Section 3.7. Section 3.8 concludes.

[5] See also Sjaastad and Scacciavillani (1996).

3.2 Review of literature

The determination of real exchange rates has always been a topic of strong interest among academics and policymakers. For economies that are largely open, as small developing countries are, the real exchange rate is indeed a key economic variable. For instance, it can be used to assess easily the competitive position of a country; it helps to detect undesirable distortions in the factor allocation such as in the Dutch disease; it can also play an important role in the emergence of large current account imbalances. For a recent discussion on the importance of real exchange rates, see Chinn (2006).

Purchasing power parity (PPP) is the basic model of exchange rate determination. It states, in its weak version, that differentials of inflation are neutralized by a corresponding adjustment of the nominal exchange rate. Therefore, if PPP holds, shocks to real exchange rates should be transitory and real exchange rates should revert over time to their long-run mean. If, on the contrary, shocks are permanent, PPP is rejected. Despite its theoretical appeal and a voluminous empirical literature, the empirical support to PPP is mixed (see Rogoff (1996) for a survey) and the idea that equilibrium exchange rates are non-stationary is now largely admitted. Variations in equilibrium real exchange rates are mostly attributed to cumulated current account imbalances, government spending, real interest rate differentials, sectoral productivity shocks (the "Balassa-Samuelson" effect), natural resources discovery, and terms-of-trade shocks.

For commodity producing countries, terms-of-trade shocks induced by changes in world commodity prices constitute a potential key determinant of their equilibrium real exchange rates. This is particularly true for small developing countries, where primary commodities represent a large share of their exports. This is also relevant for some advanced economies where the weight of the commodities in the exports is large.

Regarding advanced economies, the literature is not abundant, in the sense that advanced economies are mostly commodity importers (rather than exporters). Most studies focus on how the price of commodity imports (especially oil) affect the real exchange rate (see for examples Chen and Chen (2007) and van Amano and Norden (1998)). As regards the impact of the price of commodity exports, Chen and Rogoff (2003) study Canada, Australia and New-Zealand, and show that commodity prices have a positive influence on the real effective exchange rate of these countries. They find that the impact is stronger for New Zealand and Australia. They also suggest an analysis of which factors could determine the different countries sensibilities to commodity prices. This is precisely to question we address in Chapter 4.

Contrary to Chen and Rogoff (2003), Cashin, Cespedes, and Sahay (2004) focus their analysis on developing countries. They search for a long-run equilibrium relationship between real commodity prices and the real effective exchange rate of 58 commodity-exporting countries and find evidence of such a relationship for about one-third of their sample of countries.

Following the research initiated by Chen and Rogoff (2003) and Cashin, Cespedes, and Sahay (2004), several recent papers have provided new evidence about the long-run impact of commodity prices on real exchange rates. For instance, Koranchelian (2005) examines the relationship between oil prices and the real exchange rate of Algeria. Habib and Kalamova (2007) investigate the same relationship but for several oil producing countries. Coudert, Couharde, and Mignon (2008) consider a large set of countries, one of their objectives being to see whether the long-run impact of commodity prices on real exchange rates differs between oil exporters and non-oil commodity exporters[6]. In the most recent papers, for instance Coudert, Couharde, and Mignon (2008) or Choi and Chue (2007), panel cointegration techniques are used, whereas the results of Chen and Rogoff (2003) and Cashin, Cespedes, and Sahay (2004) were obtained with time-series techniques.

Our analysis of the long-run relationship between real exchange rates and commodity prices differs from previous ones in two respects. First, for the reasons explained in the introduction, the real exchange rate of every country of our sample is related to the price of the commodity that dominates the exports of the country. Second, cointegration relationships are tested and estimated using non-stationay panel methods robust to cross-sectional dependence.

3.3 Model

Our model consists of a simple univariate relationship between real exchange rates and commodity prices. Formally, we have:

$$REER_{i,t} = \alpha_i + \beta COM_{i,t} + \epsilon_{i,t} \tag{3.1}$$

where $REER_{i,t}$ is the real effective exchange rate (in logarithm) of country i, $COM_{i,t}$ the price of the leading export commodity (in logarithm) of country i and the error term $\epsilon_{i,t}$ is i.i.d. over periods but correlated across cross-sectional units. As in Chen and Rogoff (2003) or Cashin, Cespedes, and Sahay (2004), our model only includes a single regressor. This is motivated by the fact that several traditional explanators of real exchange rates are considered as being irrelevant for small developing countries. For instance, given that many small developing countries are poorly integrated to world financial markets, it is very

[6]See also Chen and Chen (2007) for the relation between oil and importing-countries real exchange rates.

unlikely that real interest rate differentials or net foreign asset accumulation be a significant determinant of real exchange rates. Furthermore, should some of these other determinants be relevant, like the Balassa-Samuelson effect, the data are very often not available or of low quality. In any case, the non-stationary panel approach that we use guarantees that our results be at least consistent.

Relationship 3.1 will be estimated using panel cointegration methods with monthly data covering the period 1980-2009.

3.4 Data

We use three datasets. Commodity exports come from the UN Comtrade database. These data are necessary to make the selection of countries that are specialized in the export of one particular commodity. Commodity prices are taken from the IMF International Financial Statistics database. The CPI-based real effective exchange rates come from the IMF Information Notice System database.

3.4.1 Selection of the Relevant Commodity-Country Pairs

In order to detect the countries and commodities that are eligible for our analysis, we recovered from the UN Comtrade database the annual US$ export value of 42 commodities for 68 countries over 21 years (1988-2008). Our sample of commodities is similar to the one of Cashin, Cespedes, and Sahay (2004), with the addition of oil. The list of commodities is reported in Table 3.1. The countries included in our sample are the developing and emerging countries selected by Cashin, Cespedes, and Sahay (2004), on the basis of the IMF classification of commodity countries. No advanced countries are included in our sample. The list of countries is reported in Table 3.2.

As explained in the previous sections, what is new in our approach is that we relate the real exchange rate of a particular country to the price of the commodity which is the dominant export of that country. To do so, we have to identify from our set of countries and commodities a series of country-commodity pairs. We proceed to the construction of that series as follows:

1. For every country in the sample, we compute the 1988-2008 average ratio of the export value of each commodity exported by the country in the total value of all commodity exports. For example, in the case of Mali, the average share of its exports of cotton in its total exports of all commodities is 33%.

Table 3.1: List of commodities (with their HS1992 codes)

Code	Commodity
H0-0201	Meat of bovine animals, fresh or chilled
H0-03	Fish, crustaceans, molluscs, aquatic invertebrates ne
H0-0306	Crustaceans
H0-0803	Bananas, including plantains, fresh or dried
H0-0901	Coffee, coffee husks and skins and coffee substitutes
H0-0902	Tea
H0-1001	Wheat and meslin
H0-1005	Maize (corn)
H0-1006	Rice
H0-120710	Palm nuts and kernels
H0-120810	Soya bean flour or meal
H0-1513	Coconut, palm kernel, babassu oil, fractions, refined
H0-17	Sugars and sugar confectionery
H0-18	Cocoa and cocoa preparations
H0-1801	Cocoa beans, whole or broken, raw or roasted
H0-24	Tobacco and manufactured tobacco substitutes
H0-2401	Tobacco unmanufactured, tobacco refuse
H0-2510	Natural phosphates (calcium, calcium-aluminium), chal
H0-2612	Uranium or thorium ores and concentrates
H0-2701	Coal, briquettes, ovoids etc, made from coal
H0-2704	Retort carbon, coke or semi-coke of coal, lignite,pea
H0-2705	Coal gas, water gas, etc. (not gaseous hydrocarbons)
H0-2709	Petroleum oils, oils from bituminous minerals, crude
H0-2835	Phosphatic compounds
H0-2919	Phosphoric esters, their salts and derivatives
H0-40	Rubber and articles thereof
H0-4001	Natural rubber and gums, in primary form, plates, etc
H0-52	Cotton
H0-5201	Cotton, not carded or combed
H0-7108	Gold, unwrought, semi-manufactured, powder form
H0-72	Iron and steel
H0-7201	Pig iron and spiegeleisen in primary forms
H0-74	Copper and articles thereof
H0-7401	Copper mattes, cement copper (precipitated copper)
H0-75	Nickel and articles thereof
H0-7502	Unwrought nickel
H0-76	Aluminium and articles thereof
H0-7601	Unwrought aluminium
H0-79	Zinc and articles thereof
H0-7901	Unwrought zinc
H0-80	Tin and articles thereof
H0-8001	Unwrought tin

*Notes.*The codes are based on the HS1992 classification used in the UN Comtrade database.

Table 3.2: List of countries

Countries		
Algeria	India	Papua New Guinea
Argentina	Indonesia	Paraguay
Bahrain	Iran	Peru
Bangladesh	Jordan	Qatar
Benin	Kenya	Saudi Arabia
Bolivia	Kuwait	South Africa
Brazil	Madagascar	Sri Lanka
Burundi	Malawi	Sudan
Côte d'Ivoire	Malaysia	Suriname
Cameroon	Mali	Syria
Central African Rep.	Mauritania	Tanzania
Chile	Mexico	Thailand
Colombia	Morocco	Togo
Costa Rica	Mozambique	Tunisia
Dominica	Myanmar	Turkey
Dominican Rep.	Nicaragua	Uganda
Ecuador	Niger	United Arab Emirates
Egypt	Nigeria	Uruguay
Ethiopia	Norway	Venezuela
Gabon	Oman	Yemen
Ghana	Pakistan	Zambia
Guatemala	Papua New Guinea	Zimbabwe
Honduras	Paraguay	

2. Using these country specific commodity export ratios, we then proceed to a cross-country comparison to determine, for each commodity, which country has the highest export ratio. We then keep the country that is so identified to create one particular commodity-country pair. For example, this procedure shows that Mali is the country, among all the countries of our sample, that has the highest ratio of cotton exports. We applied this procedure to the 42 commodities of our sample and we thus obtained a series of 42 commodity-country pairs.

3. As we are only concerned with countries that are highly specialized in the export of one leading commodity, we only keep for future analysis the commodity-country pairs where the commodity share is higher than 20% of total commodity exports. This threshold is set arbitrarily but other thresholds will be used as robustness checks.

4. Among the remaining pairs, it appeared that, in a few cases, the export share of the dominant commodity was very volatile over time. We then decided to eliminate all country-commodity pairs where the share of the dominant export was less than 2% during 5 years or more. We removed for instance Mozambique, because the share of its aluminum exports in total exports was more than 50% after 2001 but about zero before 2000.

In Table 3.3, we report for 21 commodities the 3 countries with the largest commodity share. Commodities not having a share of 5% in at least one country are not reported in the Table. Following the selection procedure just described, 11 commodity-country pairs are kept for our empirical analysis. They are listed in Table 3.4. The pair Cotton-Benin is not reported due to the unavailability of real exchange rate data over the period 1980-2009. We also report in Table 3.5 the structure of the commodity exports for the 11 countries of our final sample. In the last section, we do perform robustness checks using alternative samples.

An important feature about the construction of our final series of commodity-country pairs is that only one country is associated to each commodity. This implies obviously that some countries which are highly specialized in the export of a particular commodity, as for instance the OPEC countries, are not included in our final sample. But, by proceeding like this, we eliminate the risk that our results be specific to a particular commodity, for instance oil, that would be overrepresented in the final sample.

3.4.2 Commodity Prices

Commodity prices are taken from the International Financial Statistics (IFS) database of the IMF. Two series, Tobacco and Gold, were not available in IFS and were taken from Datastream (with respective codes :USI76M.ZA and GOLDBLN). We provide details of each series in Table 3.6. All the series are available at the monthly frequency over the period 1980-2009. To the extent we work on long-run relationship, we do not deseasonalize our series.

In order to capture properly the relationship between real exchange rates and commodity prices, commodity prices have to be expressed in real terms. Following Cashin, Cespedes, and Sahay (2004), we compute the real price of each commodity by deflating the price of each commodity by the IMF's index (of the unit value) of manufactured exports (MUV)[7]. The use of the MUV index as a deflator is common in the commodity-price literature (see for example Deaton and Miller (1996) and Cashin, Cespedes, and Sahay (2004)). The MUV index is represented in Figure 3.1. Real commodity prices are then normalized with base period January 1995=100. Normalized real commodity prices are represented in Figure 3.2.

We can see in Figure 3.2 that real commodity prices (especially gold, oil, uranium and copper) wandered around a long run average, or slightly decreased over time with a sudden and steep boost during last years. Banana on the contrary exhibits large fluctuations, with some potential seasonality pattern.

[7]MUV is the unit value index (in US dollars) of manufacturing exports from 20 developed countries with country weights based on the countries' total 1995 exports of manufactures (base 1995=100). This MUV index deflator is provided by the IMF's IFS database.

Table 3.3: Country's specialization in commodity exports

	Commodity weights in country's total exports		
Commodity	1	2	3
Oil	Nigeria 95.12%	Yemen 82.76%	Iran 79.72%
Cotton	Benin 61.00%	Mali 33.48%	Pakistan 20.52%
Tobacco	Malawi 60.50%	Zimbabwe 19.53%	Tanzania 6.35%
Copper	Zambia 59.99%	Chile 30.79%	Peru 13.93%
Gold	Mali 54.05%	Burundi 35.45%	Ghana 28.56%
Coffee	Burundi 50.98%	Ethiopia 46.43%	Uganda 36.87%
Uranium	Niger 41.73%	Benin 29.90%	Brazil 0.01%
Cocoa	Côte d'Ivoire 34.10%	Ghana 33.16%	Cameroon 9.75%
Aluminium	Mozambique 33.44%	Bahrain 12.89%	United Arab Emirates 12.77%
Soya	Paraguay 32.72%	Argentina 4.45%	Brazil 4.22%
Fish	Mauritania 30.96%	Mozambique 19.87%	Madagascar 14.34%
Bananas	Dominica 29.20%	Ecuador 17.83%	Costa Rica 12.83%
Tea	Kenya 21.20%	Sri Lanka 15.12%	Malawi 7.97%
Crustaceans	Mozambique 18.96%	Madagascar 13.35%	Nicaragua 10.67%
Iron	South Africa 11.24%	Zimbabwe 10.45%	Dominican Rep. 10.17%
Sugar	Guatemala 9.71%	Malawi 9.15%	Nicaragua 6.28%
Rice	Myanmar 8.99%	Uruguay 7.72%	Pakistan 7.08%
Coal	Colombia 8.85%	South Africa 6.24%	Indonesia 3.01%
Beef	Nicaragua 6.15%	Uruguay 4.78%	Paraguay 4.70%
Tin	Bolivia 6.03%	Peru 0.68%	Indonesia 0.64%
Rubber	Sri Lanka 5.81%	Thailand 5.10%	Indonesia 3.81%

*Notes.*Weights are defined as the ratio between the commodity exports of the country and all the commodity exports of the country. Annual averages over the period 1988-2008

Table 3.4: Final set of country-commodity pairs

	Commodity	Country	Weight
1	Oil	Nigeria	95.1%
2	Tobacco	Malawi	60.5%
3	Copper	Zambia	60.0%
4	Gold	Mali	54.1%
5	Coffee	Burundi	51.0%
6	Uranium	Niger	41.7%
7	Cocoa	Côte d'Ivoire	34.1%
8	Soya	Paraguay	32.7%
9	Fish	Mauritania	31.0%
10	Bananas	Dominica	29.2%
11	Tea	Kenya	21.2%
12	Crustaceans	Mozambique	19.1%
13	Aluminium	Bahrain	12.9%
14	Iron	South Africa	11.2%

Notes. Weights are defined as the ratio between the commodity exports of the country and all the commodity exports of the country. Annual averages over the period 1988-2008. In the table, we have: 5 pairs for which the main commodity has a share larger than 50%, 6 pairs larger than 40%, 9 pairs larger than 30%, 11 pairs larger than 10% and 14 pairs larger than 10%.

Table 3.5: Main commodities exported and their shares in the exports

Country	Commodity Exports		
Nigeria	Oil	Cocoa	Rubber
	95.1%	0.4%	0.3%
Malawi	Tobacco	Sugar	Tea
	60.5%	9.2%	8.0%
Zambia	Copper	Cotton	Sugar
	60%	3.2%	2.2%
Mali	Gold	Cotton	Oil
	54.1%	33.5%	1%
Burundi	Coffee	Gold	Tea
	51%	35.5%	2.5%
Niger	Uranium	Gold	Tobacco
	41.7%	8.5%	6.3%
Côte d'Ivoire	Cocoa	Oil	Coffee
	34.1%	15%	3.2%
Paraguay	Soya	Cotton	Bovin meat
	32.7%	11.2%	5.3%
Mauritania	Fish	Gold	Iron
	31%	12%	10%
Dominica	Bananas	Tobacco	Fish
	29.2%	1%	1%
Kenya	Tea	Oil	Coffee
	21.2%	10.6%	5.7%

Table 3.6: Description of the commodity price series

Commodity	Source	Description
Aluminium	IFS	Aluminum, LME standard grade, minimum purity, cif UK US$ per Metric Ton
Bananas	IFS	Central American and Ecuador, first class quality tropical pack, Chiquita, Dole and Del Monte, U.S. importer's price FOB U.S. ports (Sopisco News, Guayaquil). $/Mt
Cocoa beans	IFS	International Cocoa Organization cash price. Average of the three nearest active futures trading months in the New York Cocoa Exchange at noon and the London Terminal market at closing time, CIF U.S. and European ports (The Financial Times, London). $/Mt
Coffee (other milds)	IFS	International Coffee Organization, Other Mild Arabicas New York cash price. Average of El Salvador central standard, Guatemala prime washed and Mexico prime washed, prompt shipment, ex-dock New York. Cts/lb
Copper	IFS	London Metal Exchange, grade A cathodes, spot price, CIF European ports (Wall Street Journal, New York and Metals Week, New York). Prior to July 1986, higher grade, wirebars, or cathodes. $/Mt
Crustaceans	IFS	Shrimp, U.S., frozen 26/30 count, wholesale NY US$ per pound
Fish	IFS	Fresh Norwegian Salmon, farm bred, export price (NorStat). US$/kg
Gold	DS	Gold Bullion LBM US$/Troy Ounce
Iron	IFS	Iron Ore Carajas US cents per Dry Metric Ton Unit
Oil	IFS	U.S., West Texas Intermediate 40o API, spot price, FOB Midland Texas (New York Mercantile Exchange, New York). (In 1983-1984 Platt's Oilgram Price Report, New York). $/bbl
Soya	IFS	Soybean U.S., cif Rotterdam US$ per Metric Ton
Tea	IFS	Mombasa auction price for best PF1, Kenyan Tea. Replaces London auction price beginning July 1998. Cts/Kg
Tobacco	DS	Tobacco, US (all markets), mid month curn
Uranium	IFS	Metal Bulletin Nuexco Exchange Uranium (U3O8 restricted) price. $/lb

Notes."IFS" refers to International Financial Statistics from the IMF. "DS" refers to Datastream. "LME" refers to London Metal Exchange. "Cif" refers to cost, insurance and freight. "FOB" refers to free on board. "bbl" refers to barrel (42 US Gallons). "API" refers to American Petroleum Institute.

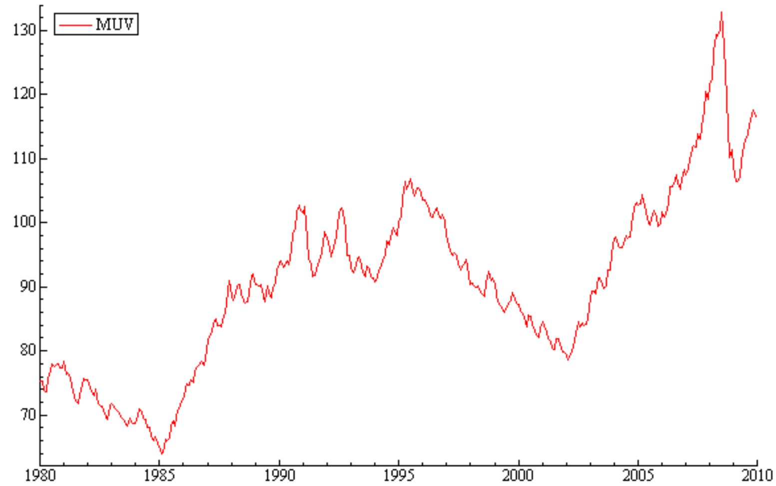


Figure 3.1: MUV: Unit Value Index (in US dollars) of manufactures (commodity deflator) - normalized version (1995=100)

3.4.3 Real Exchange Rates

As in common in many studies on the determination of real exchange rate, real exchange rates are real effective exchange rates based on consumer prices[8]. The data come from the IMF's Information Notice System (INS)[9]. To the extent we work on long-run relationship, we do not deseasonalize our series.

We represent the CPI-based real effective exchange rates (normalized to January 1995=100) in Figure 3.3 for the period 1980 to 2009. We can see that some countries experienced sudden depreciation. For example, Zambia experienced some difficult spells during the middle of the eighties when the Zambian currency fell abruptly in the context of unsuccessful IMF interventions. In addition, three countries in our selection are part of the CFA Franc zone, whose currency was devaluated by 50% against the French Franc in January 1994. Note also that the CFA Franc is linked to the euro since 1999.

To anticipate the econometric analysis, the real exchange rate and the real commodity price for our sample of 11 country-commodity pairs are plotted in Figures 3.4 and 3.5. One can observe that some series seem to be clearly correlated, while others are not. Formal statistical tests are performed later.

[8]An alternative real exchange rate, based on unit labour cost, is available only for a few countries of our sample.

[9]See Desruelle and Zanello (1997) for details regarding the construction of the real effective exchange rates.

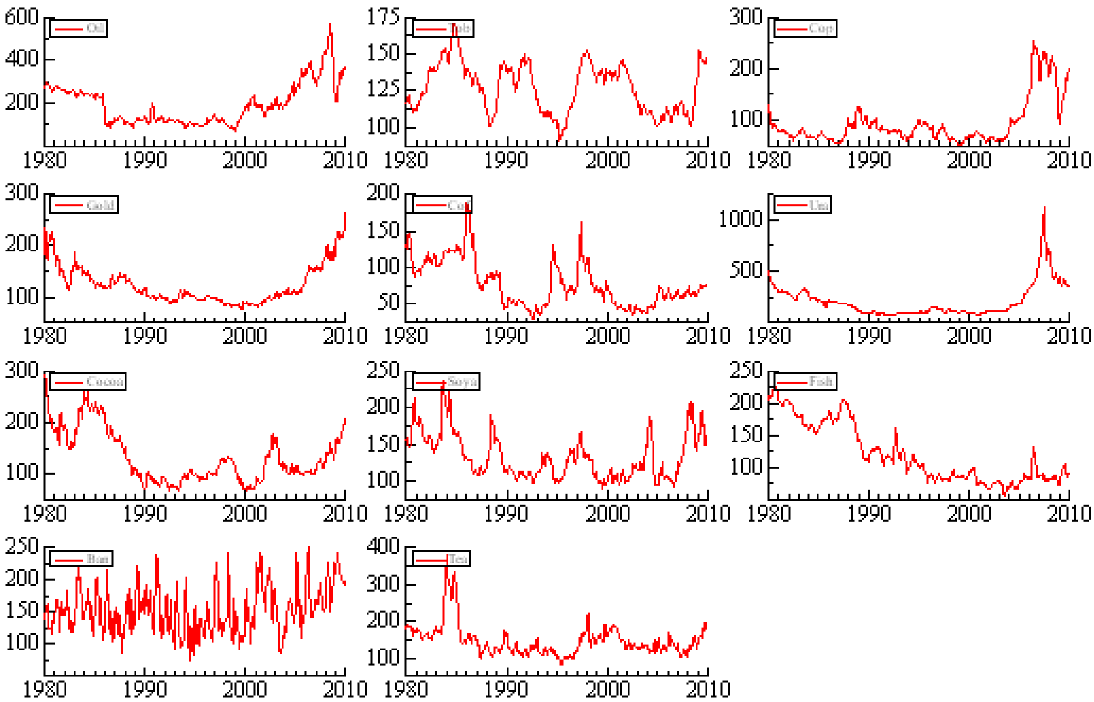


Figure 3.2: Selected commodity prices series, normalized (1995=100) and deflated (by the MUV) - by row, from left to right: oil, tobacco, copper, gold, coffee, uranium, cocoa, soybean, fish, banana and tea.

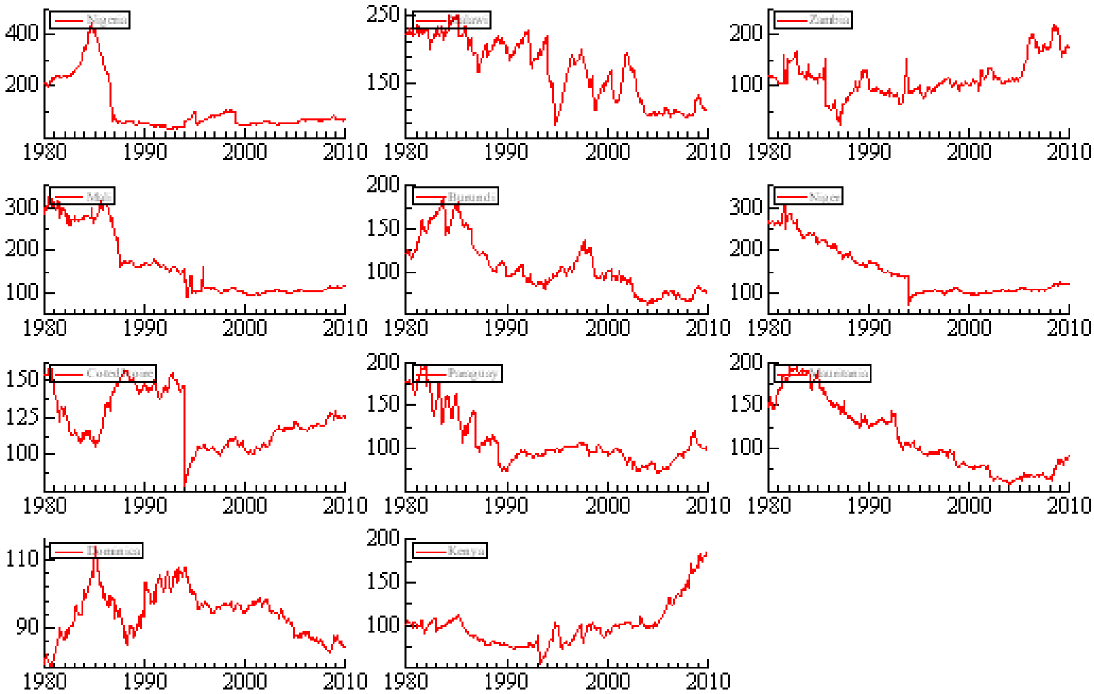
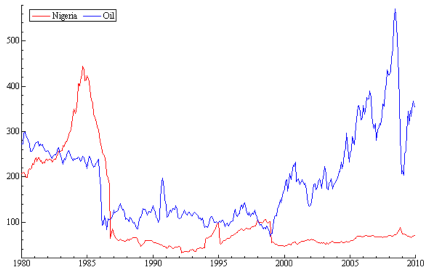
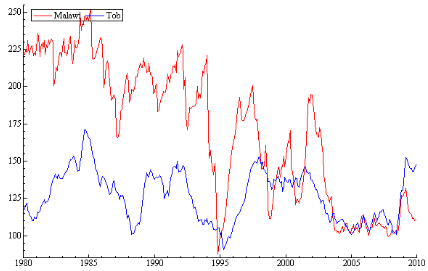


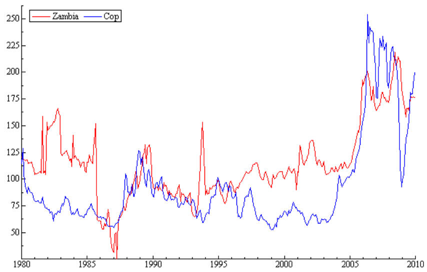
Figure 3.3: Normalized (1995=100) Real Effective Exchange Rates - by row, from left to right: Nigeria, Malawi, Zambia, Mali, Burundi, Niger, Côte d'Ivoire, Paraguay, Mauritania, Dominica and Kenya



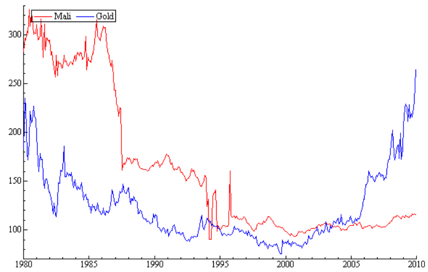
(a) Nigeria and oil



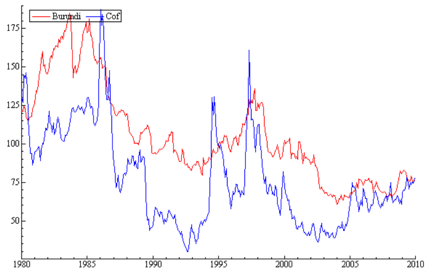
(b) Malawi and tobacco



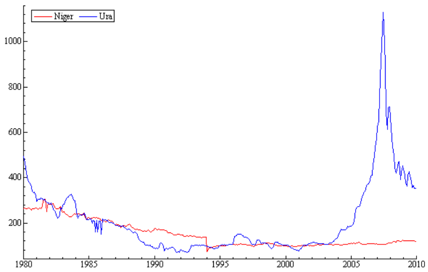
(c) Zambia and copper



(d) Mali and gold



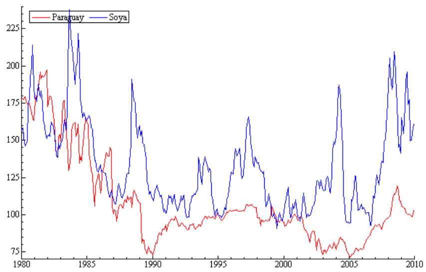
(e) Burundi and coffee



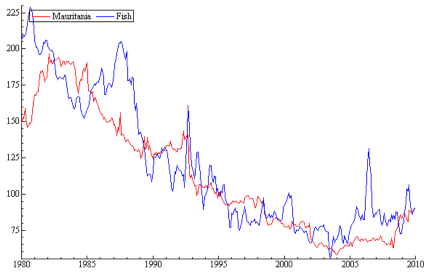
(f) Niger and uranium

Figure 3.4: Real exchange rates (dashed lines) and commodity prices (solid lines). Both series are normalized (1995=100) and commodity price are deflated by the manufactures unit value index (MUV)

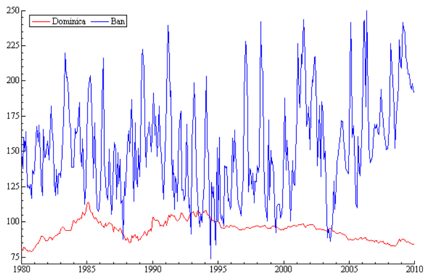
(a) Côte d'Ivoire and cocoa



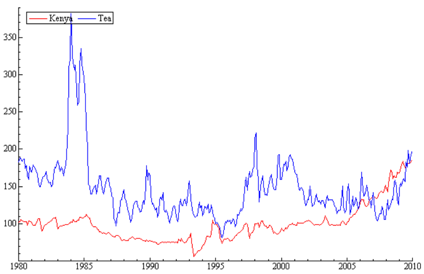
(b) Paraguay and soya



(c) Mauritania and fish



(d) Dominica and banana



(e) Kenya and tea

Figure 3.5: Real exchange rates (dashed lines) and commodity prices (solid lines). Both series are normalized (1995=100) and commodity price are deflated by the manufactures unit value index (MUV)

3.4.4 Descriptive Statistics

Some elementary statistics relative to the real exchange rates and the real commodity prices (both expressed in logs) are reported in Table 3.7. We see that there are big differences in the standard deviation across commodity prices. For example, the standard deviation of tobacco logarithmic prices is more than four times larger than the one of uranium. We may also observe differences in the variability of real exchange rate across countries, the standard deviation of the logarithmic real exchange rate of Nigeria being more than four times larger than the one of Dominica. We find unit roots for most of the series (excepted Malawi exchange rate and soya prices) using Dickey-Fuller tests (with constant and trend, or just a constant if the trend is not significant).

Table 3.7: Basic descriptive statistics

Series	min	mean	max	std.dev	ADF(p)	p
Nigeria	3.4853	4.4420	6.0982	0.67616	-1.405	0
Oil	4.2296	5.1581	6.3467	0.46921	-2.367	1
Malawi	4.4747	5.0985	5.5274	0.29364	-5.114**	8
Tobacco	4.5063	4.8337	5.1414	0.13416	-2.381	1
Zambia	3.241	4.6927	5.3882	0.32063	-3.234	1
Copper	3.9575	4.4567	5.5370	0.36605	-1.768	8
Mali	4.5016	4.9992	5.7875	0.41195	-0.7598	7
Gold	4.3250	4.7750	5.5775	0.26858	-0.5283	2
Burundi	4.1084	4.6163	5.2196	0.28290	-2.372	0
Coffee	3.3920	4.2721	5.2328	0.39297	-2.720	1
Niger	4.2787	4.9497	5.7179	0.35688	-1.073	2
Uranium	4.2433	5.1339	7.0295	0.61418	-1.612	10
Cote d'Ivôire	4.3603	4.8025	5.0642	0.14611	-2.650	0
Cocoa	4.2040	4.8177	5.6795	0.35908	-1.082	2
Paraguay	4.2700	4.6525	5.2845	0.24802	-2.359	1
Soya	4.5052	4.8685	5.4715	0.21680	-3.598*	1
Mauritania	4.0602	4.6604	5.2820	0.36381	-0.9812	0
Fish	4.0091	4.7237	5.4329	0.36617	-2.545	4
Dominica	4.3684	4.5454	4.7373	0.073745	-3.315	0
Banana	4.2995	5.0015	5.5210	0.23508	-2.995	9
Kenya	4.0221	4.5774	5.2205	0.22288	-1.111	4
Tea	4.4067	4.9619	5.9453	0.24721	-3.069	2

Notes. Series are in logarithm, normalized and, as regards commodity prices, deflated by manufacture unit value index (MUV). Critical values for the Augmented Dickey-Fuller tests (with constant and trend) are -3.42 for 5% and -3.99 for 1%.

3.5 Econometric methodology

In the literature, due to the potential non-stationarity of the series, the most common method for analyzing the dependence of the real exchange rate of highly specialized countries on the international price of their dominant commodity is related to cointegration methods (see for example Cashin, Cespedes, and Sahay (2004) who investigated a similar problem but with country-specific commodity indices). Although time-series cointegration methods could be used to test the relationship between real exchange rates and real commodity prices, as in Chen and Rogoff (2003) and Cashin, Cespedes, and Sahay (2004), we use instead panel-based cointegration techniques to conduct our econometric analysis. The panel approach is indeed a very convenient framework to explore the issues raised by our paper. One important advantage is that it enables us to estimate in just one step how the degree of export specialization affects the commodity price elasticity of the real exchange rate. In a time-series setup, a two step estimation strategy would be required to explore this issue and it is well known that the statistical inference in the second step is not trivial, which is why one step approaches are recommended[10]. Another advantage of using panel approach is that it guarantees that the unit-root based tests reach a good power.

Techniques combining panel and non-stationary series give rise to three kinds of methods and tests: panel unit root tests, panel cointegration tests and cointegration estimation and inference. One can distinguish two generations of panel methods. In the first generation, the methods are based on the assumption that panel units are cross-sectionally independent. Regarding our dataset, this assumption amounts to consider, for instance, that the real exchange rates of Mali and Kenya are independent, as would be cocoa and coffee prices. It is obvious that such an assumption is unrealistic for effective exchange rates and commodity prices, as it is confirmed formally by the results of the cross-sectional dependence test of Pesaran (2004) reported in Table 3.8. We therefore use second generation panel techniques, which allow for cross-sectional dependence. We briefly describe these methods in the following Subsections[11].

3.5.1 Panel Unit Roots Tests

In the empirical analysis, we first test for the non-stationarity of real exchange rates and of commodity prices. First generation panel unit root tests proposed by Levin, Lin, and James Chu (2002) (LLC hereafter) and Im, Pesaran, and Shin (2003) (IPS hereafter) are the most popular tests in empirical studies. LLC pool the panel series, correct and standardize the t-stat to make it normally distributed. IPS on the other hand do not pool the data but estimate separate augmented Dickey-Fuller unit root tests for cross-section units and average the t-tests. After standardization, this average follows a normal distribution. Though these tests are widely applied, it has been shown that they are inconsistent in the presence of cross-sectional dependence, as well

[10] The two step estimation strategy would comprise a country by country estimation of the cointegration coefficients in the first step and a cross-country regression analysis between the estimated cointegration coefficients and the degree of commodity export specialization in the second step.

[11] See the following surveys for a more detailed review of these techniques: Baltagi and Kao (2000), Breitung and Pesaran (2005), Hurlin and Mignon (2006), Hurlin and Mignon (2007) and Bai, Kao, and Ng (2009) for the inference technique.

Table 3.8: Pesaran test of cross-sectional dependence for different panels of countries

# countries	Pesaran Test	p-value	Dependence?	AACR
5	22.432	0.00	Yes	0.432
6	31.923	0.00	Yes	0.541
9	39.669	0.00	Yes	0.492
11	24.972	0.00	Yes	0.450
14	36.590	0.00	Yes	0.485

Notes. AACR stands for Average Absolute value of the off-diagonal elements of the Correlation matrix of Residuals. The Pesaran test and the AACR are based on fixed effects models. Results from random models are equivalent (non reported here). The null of the Pesaran test is the absence of cross sectional dependence. The samples of 5, 6, 9, 11 and 14 units are based on selection of commodity country pairs for which the weight of the dominant commodity in the country total export is the largest.

as when N (the cross-sectional dimension) is small with respect to T (the time dimension). Several alternatives and modifications of these tests have been proposed recently (see Hurlin and Mignon (2007), Breitung and Pesaran (2005) and Gengenbach, Palm, and Urbain (2010) for recent reviews of panel unit toot tests).

The second generation technique that we use is based on the subsampling approach, proposed by Choi and Chue (2007). This approach is applied to LLC and IPS and provides results robust to cross-section dependence. The idea of this approach is to approximate the limiting distribution of the tests by computing tests with smaller blocks of consecutively observed time series and to compute the empirical distribution. Hence, this method does not require estimation of nuisance parameters. As the determination of the block size may be based on two rules (stochastic calibration or minimum volatility), we will present results for both methods. In addition to the LLC and IPS tests, we also present results for an alternative test, labeled as "inverse normal panel unit root test" (INVN) (Choi (2001)), which corresponds to a generalized least squares version of the ADF test.

3.5.2 Panel Cointegration Tests

If panel unit root tests do not reject the hypothesis that real exchange rates and commodity prices are non-stationary, the next step is to check whether the series are cointegrated. If it is the case, this means that real exchange rates of highly specialized countries are tied to commodity prices of their dominant commodity. In a panel context, the usual tests are those developed by Kao (1999) and Pedroni (1999). These tests belong to the residuals based cointegration tests family. These techniques rely on the assumption of cross-sectional independence.

In order to obtain results that are robust to the cross-sectional dependence, we follow the method proposed by Fachin (2007), which consists in using a block-bootstrap version of the group and the mean t-statistic of Pedroni (1999). His method relies on fast-double bootstrap procedures of Davidson and MacKinnon (2000) which combine good size and power properties with reasonable computing power requirements. The theoretical validity of bootstrap procedures in the non-stationary panel context have recently been developed by Palm, Smeekes, and Urbain (2011).

3.5.3 Panel Cointegration Estimate and Inference

If real exchange rates and commodity prices are found to be non-stationary, the next step is to estimate the cointegration coefficient β in equation 3.1.

As it is well known, the use of normal OLS techniques leads to spurious regression when the series are non-stationary and, consequently, specific panel-cointegration techniques have to be used. Phillips and Moon (2000) show that in the case of homogeneous and near-homogeneous panels, the coefficient of cointegration can be estimated by a fully modified (FM) estimator. This method is non-parametric as it employs kernel estimators of the nuisance parameters that affect the asymptotic distribution of the OLS estimator. It tackles the possible problem of endogeneity of the regressors as well as the autocorrelation of residuals. Alternatively, Kao and Chiang (1999) and Mark and Sul (2003) propose a dynamic least square estimator (DOLS). This estimation procedure is parametric and has the advantage of computing convenience.

Though largely used in the empirical literature, these techniques have a major weakness since they assume cross-section independence. Because of this, we use the technique recently developed by Bai, Kao, and Ng (2009) (BKN), which is robust to cross-sectional dependence. They consider the following framework:

$$REER_{i,t} = \beta COM_{i,t} + e_{i,t} \tag{3.2}$$
$$COM_{i,t} = COM_{i,t-1} + \epsilon_{i,t} \tag{3.3}$$
$$e_{i,t} = \lambda_i' F_t + u_{i,t} \tag{3.4}$$
$$F_t = F_{t-1} + \eta_{i,t}, \tag{3.5}$$

where F_t are unobserved factors and λ_i the factor loadings. The BKN method imposes a factor structure on $e_{i,t}$ to capture the cross-sectional dependence. They use an iterative procedure to estimate jointly the factors and the cointegration coefficient (β). We present the results obtained with the three estimators (BKN, FMOLS and DOLS) and compare them.

3.6 Results

In this section, we successively test the non-stationarity of real exchange rates and of commodity prices, and the existence of a single homogeneous cointegration relationship between them. We then estimate the cointegration relationship.

3.6.1 Stationarity analysis

Output of the stationarity tests of commodity prices is reported in Table 3.9. We observe that each of the three tests considered (LLC, IPS and INVN) do not reject the null hypothesis of unit root in the panel. This result holds at the 10% level of significance. It is verified whether stochastic calibration or minimum volatility block selection rules are used (see Choi and Chue (2007) for details). In only one case, namely the IPS test, the null of non-stationarity is rejected at 10% (though not at 5%). Therefore, there is enough evidence to treat commodity prices as non-stationary. This result is interesting as it departs from the usual idea that commodity prices are partially predictable due to seasonality patterns (Gorton and Rouwenhorst (2006)).

Panel unit root tests for the real exchange rates are reported in Table 3.10. Without any exception, all the tests (LLC, IPS, INVN, with either SC or MV block selection rule) confirm that real exchange rates possess a panel unit root. PPP is therefore rejected.

Table 3.9: Subsampling-based panel unit root tests for commodity prices

Test	test value	block rule	crit val (5%)	block size	crit val (10%)	block size
LLC	-7.752	SC	-9.291	69	-9.064	67
		MV	-9.580	20	-8.476	33
IPS	-2.400	SC	-2.479	48	-2.347*	31
		MV	-2.489	25	-2.334*	33
INVN	0.1546	SC	-2.734	147	-2.191	176
		MV	-2.976	19	-2.500	24

Notes. Commodity prices are in logarithms. An intercept and a trend have been considered in all the experiments. "SC" and "MV" hold for Stochastic Calibration and Minimum Volatility, respectively, two alternative block selection rules. The minimum and maximum block sizes for the MV rule are 0.4 and 0.6, respectively. An asterisk indicates the rejection of the null of panel unit root. The 5% and 10% value columns give the non-centered critical values.

3.6.2 Cointegration analysis

Results of Fachin (2007)'s panel cointegration tests are reported in Table 3.11. We report both mean and median t-tests, based on Pedroni (1999), with and

Table 3.10: Subsampling-based panel unit root tests for the real exchange rates

Test	test value	block rule	crit val (5%)	block size	crit val (10%)	block size
LLC	-6.886	SC	-9.365	69	-8.064	67
		MV	-11.580	14	-8.950	33
IPS	-2.196	SC	-2.668	47	-2.494	31
		MV	-2.726	22	-2.490	33
INVN	-0.3647	SC	-3.138	147	-3.126	176
		MV	- 2.932	33	-3.061	26

Notes. Real exchange rates are in logarithms. An intercept and a trend have been considered in all the experiments. "SC" and "MV" hold for Stochastic Calibration and Minimum Volatility, respectively, two alternative block selection rules. The minimum and maximum block sizes for the MV rule are 0.4 and 0.6, respectively. An asterisk indicates the rejection of the null of panel unit root. The 5% and 10% value columns give the non-centered critical values.

without time dummies. All the tests consistently reject the null hypothesis of no cointegration at the 5 percent level. Thus, the presence of a long-run relationship between the real exchange of countries highly specialized in the export of one particular commodity and the price of that leading commodity is supported by our data. Treating each country individually, Cashin, Cespedes, and Sahay (2004) also find numerous cases where real exchange rates and commodity prices are cointegrated but, in their study, a country specific commodity price index is used rather than the price of a single commodity as it is the case here.

Table 3.11: Fachin (2007) Panel Cointegration Tests

	Test value	p-values		
		basic bootstrap	FDB1	FDB2
With common time dummies				
Mean CADF	-3.36	0	0	0
Median CADF	-3.24	1	0	-1
Without common time dummies				
Mean CADF	-3.44	0	1	0
Median CADF	-3.40	0	0	0

Notes. Real exchange rates and commodity prices are in logarithms. An intercept and a trend have been considered in all the tests. Block size selection for the cointegration test is based on 0.1T. "FDB1" and "FDB2" hold for Fast Double Boostraps of types 1 and 2 (see Fachin (2007)). P-values are in percent. FDB2 p-values can be negative, which is not due to computation errors but to the correction terms in the FDB2 formula. See Davidson and MacKinnon (2000) p.7 for comments related to this potential negativity.

3.6.3 Estimation of the cointegration relationship

Following the results of the previous subsection, we now proceed to the estimation of the cointegration relationship. We compare the results obtained with first generation techniques (DOLS and FMOLS) to the estimates provided by BKN which, as discussed before, is robust to cross-section dependence. Results are reported in Table 3.12.

Table 3.12: Cointegration esimates by DOLS, FMOLS and BKN methods

Test	β	Standard errors
Panel DOLS	0.343*	0.0021
Panel FMOLS	0.333*	0.0021
BKN	0.165*	0.00907

Notes. An intercept and a trend have been considered in all the tests. An asterisk indicates that the coefficients are significant at a level of 5 percent. Panel DOLS are based on 4 leads and lags. Panel FMOLS based on Fejer kernel with window length $3.21 * T^{1/3}$. BKN is the CupFM estimator (see Bai, Kao, and Ng (2009)). One factor was sufficient for the BKN estimate, convergence occured after 6 iterations, with quadratic spectral kernel. The estimates are made for a sample of countries where the dominant commodity has a share of at least 20% in total exports.

Whatever technique is used (DOLS, FMOLS, BKN), it turns out that the cointegration coefficient is significant and positive, as expected. The coefficient estimates obtained with DOLS and FMOLS are respectively 0.343 and 0.333, which suggests that commodity price shocks have a strong long-run impact on the real exchange rate of highly specialized countries. For example, according to these results, the real exchange rate of an oil producing country should increase by 3% when the international price of oil increases permanently by 10%. Our coefficient estimate is however smaller than the 0.6484 coefficient found by Coudert, Couharde, and Mignon (2008). They also use DOLS but their regressor is a country-specific commodity price index rather than the price of the dominant commodity.

When the BKN methodology is used, the estimate of the cointegration coefficient is 0.165, thus much smaller than the coefficient obtained with DOLS and FMOLS. By correcting the bias induced by cross-section dependence, we thus find that there is still a positive and significant long-run impact of commodity prices on the real exchange rate of highly specialized countries, but the impact is now economically small. Our results also show that the results found in earlier studies with DOLS and FMOLS estimates are overestimated.

3.7 Robustness checks

To verify whether our results are specific to the selection of countries that we made, we replicated the cointegration analysis and the BKN estimation for other selections of countries. So far, a country has been considered as having a dominant commodity export if the share of that particular commodity in its total exports was at least 20%. Following the methodology presented in Section 3.4.1, we constructed five new samples of countries by fixing the minimum export share at 50%, 40%, 30%, 15% and 10%. The samples get larger as the threshold decreases: 5 countries for 50%, 6 for 40%, 9 for 30%, 12 for 15% and 14 for 10%. They show that cointegration holds from a threshold of about 20%. Indeed, when the threshold is set at 10% and 15%, we cannot reject for all tests the null of no cointegration between the commodity price of the leading commodity and the real exchange rate (see Table 3.13.) In addition, the results in Table 3.14, which are illustrated in Figure 3.6, show that the larger the share of the leading commodity in total exports, the larger the elasticity between real exchange rates and commodity prices.

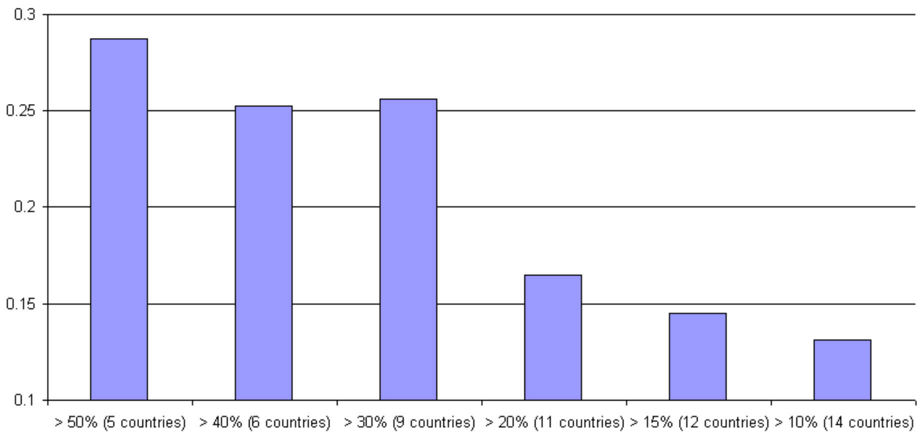


Figure 3.6: BKN cointegration coefficient for different samples (depending on the weight of the main commodity)

We found in Table 3.12 that DOLS and FMOLS methods tend to overestimate the cointegration coefficient, compared to the BKN method, which is robust to cross-section dependence. We reiterated our estimates with the samples obtained by setting the threshold at 50%, 40% and 30%, which are the cases where cointegration cannot be rejected. The results are reported in Table 3.15 and confirm that neglecting cross-sectional dependence leads to a higher elasticity estimation.

Table 3.13: Fachin (2007)'s cointegration tests for different samples

	Test	Stat	p-Btst	FDB1	FDB2	Coint.?
14 countries (> 10%)						
With common time dummies	Mean CADF	-2.88	11	9	9	No
	Median CADF	-2.76	18	13	13	No
Without common time dummies	Mean CADF	-3.22	1	1	2	yes
	Median CADF	-3.38	0	1	0	yes
12 countries (> 15%)						
With common time dummies	Mean CADF	-2.99	7	4	5	Yes
	Median CADF	-2.90	17	11	15	No
Without common time dummies	Mean CADF	-3.31	0	0	0	yes
	Median CADF	-3.38	0	0	-1	yes
11 countries (> 20%)						
With common time dummies	Mean CADF	-3.36	0	0	0	yes
	Median CADF	-3.24	1	0	-1	yes
Without common time dummies	Mean CADF	-3.44	0	1	0	yes
	Median CADF	-3.40	0	0	0	yes
9 countries (> 30%)						
With common time dummies	Mean CADF	-3.39	2	1	1	yes
	Median CADF	-3.12	3	0	-2	yes
Without common time dummies	Mean CADF	-3.59	0	0	0	yes
	Median CADF	-3.71	0	2	0	yes
6 countries (> 40%)						
With common time dummies	Mean CADF	-3.46	2	2	1	yes
	Median CADF	-3.29	6	5	6	yes
Without common time dummies	Mean CADF	-3.87	0	1	0	yes
	Median CADF	-3.82	0	1	0	yes
5 countries (> 50%)						
With common time dummies	Mean CADF	-3.57	0	0	-1	yes
	Median CADF	-3.09	16	19	18	No
Without common time dummies	Mean CADF	-3.96	0	3	0	yes
	Median CADF	-3.94	0	2	0	yes

Notes. The sample of respectively 14, 12, 11, 9, 6 and 5 countries include countries where the dominant commodity has a share of at least 10%, 15%, 20%, 30%, 40% and 50% in total exports. The 3 p-values are referred to as p-Btst, FDB1 and FDB2. Some FDB2 are negative, which is not due to computation errors but to the corrections in the FDB2 formula. See Davidson and MacKinnon (2000) p.7 for comments related to this potential negativity.

Table 3.14: Cointegration tests and estimates for different samples

Weight of the main cdty	# countries	Cointegration tests	Coefficients	SE
More than 50%	(5 countries)	C	0.2871	(0.01258)
More than 40%	(6 countries)	C	0.2528	(0.009805)
More than 30%	(9 countries)	C	0.2561	(0.008585)
More than 20%	(11 countries)	C	0.1650	(0.009072)
More than 15%	(12 countries)	NC/C	0.1448	(0.01134)
More than 10%	(14 countries)	NC	0.1312	(0.01010)

Notes. Countries in each panel are selected according to the weight of their main commodity in their total commodity exports. Cointegration test refers to mean CADF and median CADF with common time dummies of Fachin (2007). We report cointegration (C) if at least 5 of the 6 p-values are not larger than 10%. We report non-cointegration (NC) if at least 5 of the 6 p-values are larger than 10%. Otherwise, we report C/NC. The coefficients are cointegration estimates obtained with the methodology of BKN. SE holds for standard errors.

Table 3.15: Cointegration esimates by the DOLS, FMOLS and BKN methods for different samples

# countries	min weight	PDOLS	FMOLS	BKN
5	50%	0.536* (0.0031)	0.527* (0.0031)	0.2871* (0.01258)
6	40%	0.402* (0.0029)	0.395* (0.0028)	0.2528* (0.009805)
9	30%	0.360* (0.0023)	0.351* (0.0023)	0.2561* (0.008585)
11	20%	0.343* (0.0021)	0.333* (0.0021)	0.1650* (0.009072)

Notes. The different samples of respectively 5, 6, 9 and 11 countries are based on commodity country pairs where the weight of the main commodity is at least 50%, 40%, 30% and 20% of total exports. An intercept and a trend have been considered in all the tests. An asterisk indicates that the coefficients are significant at a level of 5 percent. Standard errors in parentheses

We also assessed the robustness of our results to the presence of outliers in the series of real exchange rates (resulting for instance from large devaluations). We consider as an outlier a real exchange rate monthly variations larger than 50% (see Table 3.16). Outliers are removed by simply setting the monthly variation equal to zero. We then reiterate our tests with the modified dataset. As reported in Tables 3.17 and 3.18, the results are similar to those obtained with the original dataset. To further check the potential impact of outliers, we reiterate our estimates by eliminating as above the monthly variations of the real exchange rates larger than 50%, but also the January 1994 monthly variations for Mali, Côte d'Ivoire and Niger due to the devaluation of the CFA Franc. One can see in Tables 3.17 and 3.18 that the main findings of our analysis are unchanged.

Table 3.16: List of outliers

Country	Outliers variation	Month
Nigeria	-65.3%	1986-10
Zambia	+133.7%	1987-5
	-59.2%	1985-10
	+55.1%	1987-2
	+50.4%	1981-8
Mali	+51.9%	1995-10

Notes. List of monthly variations of the real exchange rates larger than 50%.

3.8 Conclusion

The last decade has seen a sustained increase in the price of agricultural and mineral commodities. Whether commodity price booms have a positive or negative impact on the output of commodity exporting countries remains unclear so far. One particular reason why the impact could be negative is an appreciation of the real exchange rate in response to the increase of commodity prices. In this paper, we explored this relationship between real exchange rates and commodity prices for a set of developing countries specialized in the export of one leading commodity. We show that the real exchange rate appreciates when the price of the leading commodity exported by the country increases, provided that the dominant commodity accounts for at least 20 percent of the total exports of the country. We also showed that the larger the share of the main exported commodity, the stronger is the impact on the real exchange rate.

Our results therefore suggest that small developing countries heavily specialized in the export of one commodity are vulnerable to "Dutch disease" effects. One way to prevent the emergence of these effects would be to ease monetary (or

Table 3.17: Break robustness check - panel unit root tests for the real exchange rates

Removal of monthly variations larger than 50%						
Test	test value	block rule	crit val (5%)	block size	crit val (10%)	block size
LLC	-6.763	SC	-8.351	69	-7.841	67
		MV	-8.928	28	-8.077	26
IPS	-2.052	SC	-2.542	47	-2.353	31
		MV	-2.523	31	-2.341	29
INVN	0.3392	SC	-2.773	147	-2.626	176
		MV	- 2.445	31	-2.062	33
Removal of monthly variations larger than 50% and of January 1994 variations in Mali, Côte d'Ivoire and Niger						
Test	test value	block rule	crit val (5%)	block size	crit val (10%)	block size
LLC	-6.753	SC	-8.075	69	-7.460	67
		MV	-8.721	28	-8.039	26
IPS	-2.077	SC	-2.489	47	-2.343	31
		MV	-2.627	19	-2.339	30
INVN	1.272	SC	-1.510	147	-1.496	176
		MV	- 2.302	32	-2.035	31

Notes. Subsampling-based panel unit root tests. Real exchange rates are in logarithms. An intercept and a trend have been considered in all the experiments. "SC" and "MV" hold for Stochastic Calibration and Minimum Volatility, respectively, two alternative block selection rules. The minimum and maximum block size for the MV rule are 0.4 and 0.6, respectively. An asterisk indicates the rejection of the null of panel unit root. The 5% and 10% value columns give the non-centered critical values.

exchange rate) policy when there is a long and lasting increase in commodity prices.

Table 3.18: Break robustness check - Fachin (2007)'s panel cointegration tests

Removal of monthly variations larger than 50%				
	Test value	p-values		
		basic bootstrap	FDB1	FDB2
With common time dummies				
Mean CADF	-3.10	0.00	0.00	-1.00
Median CADF	-3.02	6.00	5.00	6.00
Without common time dummies				
Mean CADF	-3.23	0.00	0.00	0.00
Median CADF	-2.99	2.00	1.00	0.00
Removal of monthly variations larger than 50% and of January 1994 variations in Mali, Côte d'Ivoire and Niger				
	Test value	p-values		
		basic bootstrap	FDB1	FDB2
With common time dummies				
Mean CADF	-3.21	0.00	0.00	-1.00
Median CADF	-3.16	0.00	0.00	-4.00
Without common time dummies				
Mean CADF	-3.23	0.00	1.00	0.00
Median CADF	-2.99	3.00	2.00	3.00

Notes. Real exchange rates and commodity prices are in logarithms. An intercept and a trend have been considered in all the tests. Block size selection for the cointegration test is based on 0.1T. "FDB1" and "FDB2" hold for fast double boostraps (see Fachin (2007)). P-values are in percent. See Davidson and MacKinnon (2000) p.7 for comments related to the potential negativity of FDB2.

Chapter 4

Real Exchange Rates, Commodity Prices and Structural Factors in Developing Countries[1]

Abstract

This paper provides new empirical evidence about the relationship that may exist between real exchange rates and commodity prices in developing countries that are specialized in the export of a main primary commodity. It investigates how the exchange rate regime, the degree of financial and trade openness, the degree of export concentration and the type of the commodity exports affect the magnitude of the effects of commodity price shocks on real exchange rates.

Keywords: Real exchange rates * commodity prices * exchange rate regime * financial openness * panel analysis.

JEL Classification: C32, C33, E31, F32, 011.

[1] Joint work with Vincent Bodart and Bertrand Candelon.

4.1 Introduction

This paper provides new empirical evidence about the relationship that may exist between real exchange rates and commodity prices in developing countries that are specialized in the export of a main primary commodity. It is largely documented that for many developing countries that are dependent on the production of primary commodities, commodity price shocks may have important economic implications, either positive or negative[2]. As shown by the literature on the "Dutch disease"[3], the principal channel through which commodity price shocks may affect a country's economic performances is the real exchange rate. For that reason, the dependence of real exchange rates on commodity prices (or, more generally, on terms of trade) has been the subject of numerous empirical studies, that usually suggest that increases in the world price of commodity prices are associated with an appreciation of the real exchange rate[4]. With a very few exceptions, the purpose of these studies is strictly limited to the estimation of the real exchange rate response to commodity price shocks. However, what determines the magnitude of the real exchange rate reaction is not examined in these studies. It is the purpose of our paper to fill this gap. We do that by exploring the role played by several structural factors in shaping the real exchange rate - commodity price relationship.

Our analysis focuses on five potential determinants : the exchange rate regime, the degree of financial openness, the degree of trade openness, the degree of export diversification and the type of the main commodity exported by the country. From a policy stand point, understanding the role of these factors is particularly important given that many developing countries have been or are faced with questions such as which exchange rate regime to choose, how and by how much open the capital account, and whether and how reduce the concentration of exports on a few products. As shown in the next Section, the focus on these factors is also dictated by theoretical considerations.

Despite its narrow focus, our analysis is at the cross-section of two important topics of the literature about the determinants of economic growth in devel-

[2]For instance, the literature on the "Natural Resource Curse" suggests that increases in commodity prices have adverse, rather than positive, effects on the economic growth of commodity producing countries. For a recent survey on this topic, see Frankel (2010b)

[3]The term "Dutch disease" usually refers to the decline in the production of several sectors that is caused by a favorable shock such as a large natural resource discovery or a rise in the world market price of a primary commodity. The main source of the decline in sectoral output is an appreciation of the real exchange rate. For a nice non-technical discussion of the "Dutch disease" phenomenon, see Brahmbhatt, Canuto, and Vostroknutova (2010). For more detailed analyses, see Corden (1984) and Corden and Neary (1982).

[4]Recent studies are Chen and Rogoff (2003), Broda (2004), Cashin, Cespedes, and Sahay (2004), Coudert, Couharde, and Mignon (2008) and Bodart, Candelon, and Carpantier (2011). Coudert, Couharde, and Mignon (2008) provides a comprehensive survey of these studies.

oping countries. As reminded above, the first one is about the impact that commodity price shocks may have on the economic performances of developing countries that are specialized in the export of primary products as many small African countries are. The second topic is about the influence of structural factors such as the exchange rate regime or the degree of financial openness on economic growth.

Previous empirical evidence on what determines the strength of the real exchange rate - commodity price relationship is very scarce. To our knowledge, it is almost limited to Broda (2004) who examines whether the response of real exchange rates to terms-of-trade shocks differ systematically across exchange rate regimes[5]. He shows that in response to a fall in terms-of-trade, there is a small and slow depreciation of the real exchange rate in developing countries with a currency peg but a large and immediate real exchange rate depreciation in country where the exchange rate is floating. His analysis also concludes that the response of the real exchange rate does not differ significantly across regimes when the terms-of-trade shock is positive. The role played by structural factors in shaping the relationship between real exchange rates and commodity prices is also evoked by Chen and Rogoff (2003) who estimate such a relationship for Australia, Canada and New Zealand. They find that the relation is strong for Australia and New Zealand but less robust for Canada. They suggest that this difference is due to the fact that the Canadian dollar is de facto tied to the US dollar, while the Australia and NZ dollars are floating. A second explanation that they put forward is that commodities constitute a smaller share of the Canadian exports compared to Australia and New Zealand. Their tentative explanations are however not tested formally.

As a background material to the analysis conducted in this paper, we report in Table 4.1 estimates of the long-run (cointegrating) relationship between the real exchange rate of 33 developing countries that produce a primary commodity that counts for at least 20 percent of their total exports and the world market price of their main commodity export. The estimates of the long-run commodity price elasticity of the real exchange rate are significant for 17 countries. We can also observe that the elasticity estimates vary considerably across countries, as they range from -0,17 (Dominica Republic) to 10.39 (Ghana), with a median value of 0.21[6].

[5] An analysis similar to that of Broda (2004) has been realized recently for 9 East Asian countries by Dai and Chia (2008). Edwards (2011) and Edwards and Yeyati (2003) also examine empirically the economic impact of terms of trade disturbances under alternative exchange rate regimes, but their evidence is limited to the impact on GDP growth.

[6] To our knowledge, the only published study that provides estimates of the commodity price elasticity of the real exchange rate for a large number of countries is Cashin, Cespedes, and Sahay (2004), who report long-run (cointegrating) elasticity estimates for 19 countries. Their elasticity estimates range from 0.16 to 2.03, with a median value of 0.42

Table 4.1: Data summary

	Comdty	Country	β coint	Currency Reg			Fin Opnss			Trade Opnss			Div		Cdty Type
				1	2	3	1	2	3	1	2	3	Wght	Cat	
1	Oil	Algeria	0.39	0	27	1	28	0	0	3	25	0	46.4	2	Energy
2	Oil	Bahrain	0.09	28	0	0	0	0	28	0	0	28	23.2	1	Energy
3	Coffee	Burundi	0.65***	4	23	1	28	0	0	24	4	0	54.5	3	Agri
4	Oil	Cameroon	0.03	28	0	0	18	10	0	14	14	0	40.7	2	Energy
5	Copper	Chile	0.07	2	25	1	18	3	7	1	27	0	29.7	1	Metal
6	Oil	Colombia	0.15	0	28	0	24	4	0	28	0	0	20.6	1	Energy
7	Cocoa	Cote d'Iv.	-0.13	28	0	0	18	10	0	0	26	2	34.5	2	Agri
8	Bananas	Dominica	-0.17***	28	0	0	8	20	0	0	0	28	27.8	1	Agri
9	Oil	Ecuador	0.20	10	7	11	0	24	4	0	28	0	38.6	2	Energy
10	Coffee	Ethiopia	0.00	10	18	0	28	0	0	24	4	0	47.4	2	Agri
11	Oil	Gabon	0.13*	28	0	0	12	16	0	0	9	19	74.3	3	Energy
12	Cocoa	Ghana	10.39***	0	13	15	26	2	0	8	12	8	33.7	2	Agri
13	Coffee	Honduras	0.78	5	22	1	9	19	0	0	14	14	23.5	1	Agri
14	Oil	Iran	1.59***	0	26	2	20	8	0	18	10	0	79.7	3	Energy
15	Tea	Kenya	2.95***	7	19	2	16	0	12	0	28	0	21.4	1	Agri
16	Oil	Kuwait	0.08**	5	23	0	0	0	28	0	6	22	58.7	3	Energy
17	Tobacco	Malawi	0.72***	3	12	13	27	1	0	0	26	2	58.5	3	Agri
18	Gold	Mali	1.14***	28	0	0	12	16	0	0	28	0	52.3	3	Metal
19	Fish	Mauritania	0.52**	0	28	0	27	1	0	0	9	19	33.8	2	Agri
20	Uranium	Niger	0.20***	28	0	0	12	16	0	13	15	0	40.5	2	Metal
21	Oil	Nigeria	0.48	0	18	10	15	13	0	5	22	1	95.5	3	Energy
22	Cotton	Pakistan	1.22**	2	26	0	28	0	0	28	0	0	21.1	1	Agri
23	Oil	Papua N. G.	0.06***	10	18	0	9	19	0	0	0	28	22.9	1	Energy
24	Soya	Paraguay	0.99***	0	25	3	10	9	9	4	13	11	32.7	2	Agri
25	Oil	Qatar	0.13***	28	0	0	0	0	28	0	20	8	46.1	2	Energy
26	Oil	Saudi Ar.	0.16	28	0	0	0	0	28	0	23	5	74.9	3	Energy
27	Oil	Sudan	-0.05	8	20	0	19	9	0	25	3	0	25.4	1	Energy
28	Oil	Syria	0.38*	0	28	0	28	0	0	5	23	0	54.4	3	Energy
29	Gold	Tanzania	1.87	0	25	3	25	3	0	4	24	0	22.5	1	Metal
30	Oil	UAE	0.05	27	1	0	0	0	28	0	1	27	37.8	2	Energy
31	Coffee	Uganda	0.22	3	16	9	14	4	10	25	3	0	40.7	2	Agri
32	Oil	Venezuela	0.21	3	13	12	5	15	8	2	26	0	64.4	3	Energy
33	Copper	Zambia	0.68***	1	4	23	16	0	12	0	28	0	59.7	3	Metal
Mean			0.79										43.6		
Med.			0.21										40.5		
All				352	465	107	500	222	202	231	471	222			

Notes. Beta cointegration coefficients are computed by FMOLS in a univariate setting on annual data. *, ** and *** hold for significance at 10%, 5% and 1%. Currency regime is a de facto classification into 3 categories from 1 (fix) to 3 (flexible). Since a country currency regime can change over time, figures correspond to the number of years in each regime. Financial openness, based on Kaopen index, is classified into 3 categories from 1 (closed) to 3 (open). Since financial openness can change over time, figures correspond to the number of years in each regime. Trade openness is computed by dividing the sum of exports and imports on GDP. Trade openness is classified into 3 categories from 1 (closed) to 3 (open). Since a country can change of category over time, figures correspond to the number of years in each category. The degree of diversification, proxied by the weight (in %) of the dominant commodity in the exports, is classified into 3 categories from 1 (high diversification) to 3 (low diversification). The weight average of the dominant commodity is by definition not changing over time. Commodity type refers to the nature of the commodity, from labor low intensive (energy) to labor high intensive (agricultural).

To achieve our analysis, we use panel data covering 33 small developing countries over the period 1980-2007. Our main findings can be summarized as follows. First, we find evidence that the real exchange rate of countries specialized in the production of a main primary commodity is related in the long-run to the international price of the main commodity that they export. Second, we find evidence suggesting that the long-run commodity price elasticity of the real exchange rate varies with the exchange rate regime, the degree of trade openness, the degree of export diversification and the type of the primary commodity that is exported. Conversely, our evidence suggests that the degree of financial openness does not affect the long-run response of real exchange rates to international commodity price disturbances.

The rest of the paper is structured as follows. In Section 4.2, we outline a simple theoretical model to illustrate how the exchange rate regime, the degree of financial and trade openness, the extent of export concentration and the type of commodity exported may affect the relation between a country's real exchange rate and the world market price of its primary commodity exports. Section 4.3 describes the data while the econometric methodology and the results are presented in Section 4.4. Conclusions and policy implications are drawn in Section 4.5.

4.2 Theoretical considerations

The purpose of this Section is to illustrate with a simple theoretical framework how structural factors like the exchange rate regime or the degree of financial openness may affect the relationship between the real exchange rate a small commodity exporting country and the price of the main commodity exported by that country .

To do so, we present a model of a small open economy composed of a tradable and a non tradable sectors[7]. In addition of being pretty standard, this model has the advantage of being simple enough to be analytically tractable but rich enough to provide interesting insights. The main features of the model are as follows. It is assumed that the economy produces two goods, a primary commodity that is not consumed locally, and a nontradable good that is only available to domestic consumer. Private agents can also consume an imported consumer good. They therefore derive their utility from the consumption of the nontradable good produced domestically and the imported good. Domestic agents take the world market price of the exported commodity and of the imported consumer good as given. As our empirical analysis is only concerned

[7] The model developed in this Section is derived from the model developed by Gregorio and Wolf (1994). For a detailed presentation of small open economy models with tradable and non tradable goods, see Dornbusch (1980) and Obstfeld and Rogoff (1996)

with the long-run relationship that may exist between real exchange rates and commodity prices, our model has no dynamics.[8]

In what follow, the exported, nontradable and imported goods are denoted respectively by the suffix X, N, and M.

On the production side, it is assumed that the exported primary commodity (Y_x) is produced with a technology that combines labor (L_x) and capital (K_x). The production of the nontradable good (Y_n) however only requires labor (L_n). Labor is perfectly mobile across the two sectors but capital is specific to the exportable sector. The production function of the two goods is assumed to be Cobb-Douglas:

$$Y_x = a_x L_x^{\alpha} K_x^{1-\alpha} \tag{4.1}$$
$$Y_n = a_n L_n, \tag{4.2}$$

where a_x and a_n are exogenous productivity factors and $0 < \alpha < 1$.

As it is the case for many commodities, it is assumed that in the long run the domestic price of the exported good (P_x) is determined by the law of one price:

$$P_x = EP_x^*, \tag{4.3}$$

where E is the nominal exchange (defined as the number of units of domestic currency per one unit of the foreign currency) and P_x^* is the world market price of the commodity.

Let's denote w, the wage rate paid to labor, and r, the domestic rate of interest.

From standard profit maximization, we can derive the following expressions relating the price of each good to the price of the production factors:

$$P_x = \left(\frac{\psi_x}{a_x}\right) w^{\alpha} r^{1-\alpha} \tag{4.4}$$
$$P_n = \frac{w}{a_n}, \tag{4.5}$$

[8] As shown in details by Corden (1984), we are conscious that the conclusions delivered by small open economy models about the determination of the real exchange rate are very much dependent on several assumptions, in particular those about the number of sectors, the technology of production in each sector and the degree of factor mobility across sectors and countries. These issues are however neglected here as our purpose is simply to have a framework that can illustrate how structural factors influence the determination of the real exchange rate.

where: $\psi_x = \alpha^{-\alpha}(1-\alpha)^{-(1-\alpha)}$.

Let's now denote q, the real exchange rate as the ratio between the price of the nontradable good and the price of the exportable good. Combining Equation (4.4) and Equation (4.5), we obtain the following expression for q:

$$q = \frac{P_n}{P_x} = \left(P_x^{(1-\alpha)} a_x a_n^{-\alpha} \psi_x^{-1} r^{-(1-\alpha)}\right)^{\frac{1}{\alpha}}. \tag{4.6}$$

On the demand side, households consume two goods, the nontradable good produced domestically and an imported good. Individual preferences are given by a Cobb-Douglas utility function. Denoting C_n, the consumption of the nontradable good, and C_m, the consumption of the imported good, we have :

$$U = kC_n^{\phi} C_m^{1-\phi}, \tag{4.7}$$

where $k = \left(\phi^{\phi}(1-\phi)^{(1-\phi)}\right)^{-1}$ is a constant such $0 < \phi < 1$.

The consumer has thus constant expenditures shares ϕ and $(1-\phi)$ on the nontradable and the imported goods. Letting Z denote the consumer income, we thus have the following demand functions:

$$C_n = \phi \frac{Z}{P_n} \tag{4.8}$$

$$C_m = (1-\phi)\frac{Z}{P_m}, \tag{4.9}$$

where the domestic price of the imported good, P_m, is determined by the law of one price: $P_m = EP_m^*$, and the domestic economy takes P_m^* as given. As there is no government in our model, the consumer income is simply equal to nominal output: $Z = P_n Y_n + P_x Y_x$.

The model of the domestic economy is closed by the equilibrium conditions for the nontradable good market and the labour market, respectively:

$$C_n = Y_n \tag{4.10}$$

$$L = L_n + L_x, \tag{4.11}$$

where L, the labor supply, is fixed and exogenously given.

For later use, notice that by combining Equations (4.8) and (4.10) with the definition of Z, we can rewrite the expression for C_n and Z as follows:

$$C_n = (\frac{\phi}{1-\phi})(\frac{P_x Y_x}{P_n}) \tag{4.12}$$
$$Z = \frac{P_x Y_x}{1-\phi}. \tag{4.13}$$

Notice also that it can be shown with some simple algebra that the labor market equilibrium condition implies that the labor demand from each sector can be expressed as a constant ratio of the total labor supply. To save space, only the expression for L_x is reported here:

$$L_x = \left(\frac{\alpha(1-\phi)+\phi}{\alpha(1-\phi)}\right)^{-1} L. \tag{4.14}$$

Using the above expressions, the joint equilibrium in the market of the non-tradable good and the labor market is given by:

$$q = \alpha(\frac{a_x}{a_n})h^{(1-\alpha)}L^{(\alpha-1)}K_x^{(1-\alpha)}, \tag{4.15}$$

where $h = \frac{\alpha(1-\phi)+\phi}{\alpha(1-\phi)}$.

Regarding the foreign economy, it is assumed for simplicity that it produces only the final good M (Y_m^*), with a technology that combines labor (L_m^*) and the primary commodity (X_m)[9]:

$$Y_m^* = a_m {L_m^*}^{\gamma} X_m^{1-\gamma}, \tag{4.16}$$

with $0 < \gamma < 1$.

From the first order conditions of profit maximization, we obtain that the cost of one unit of the foreign consumer good is given by:

[9] As in Cashin, Cespedes and Sahay (2004), the foreign economy does not correspond to the rest of the world, which also includes other countries producing the primary commodity.

$$P_m^* = \left(\frac{\psi_m}{a_m}\right) w^{*\gamma} P{*_x}^{1-\gamma}, \tag{4.17}$$

where $\psi_m = \gamma^{-\gamma}(1-\gamma)^{-(1-\gamma)}$.

Finally, notice that the model can be reduced to four equations: Equation (4.6), which describes the equilibrium of the firm in the domestic economy; Equation (4.15), which represents the joint equilibrium of the labor market and the market for the nontradable good; Equation (4.17) which gives the equilibrium of the foreign producer; and Equation (4.3), which is the law of one price for the exportable good. These four equations determine: q, the real exchange rate price, P_x, the price of the primary commodity expressed in domestic currency, P_m^*, the world market price of the foreign consumer good and either E, r, or K_x depending on the assumptions that will be made regarding the exchange rate regime and the degree of international capital mobility.

So far, we have defined the real exchange rate as the relative price of the nontradable good in terms of the tradable good. As it is well known, the real exchange rate can be defined in many different ways[10] . It is for instance very common to define the real exchange rate as the ratio between the domestic consumer price index (P) and the foreign consumer price index expressed in the domestic currency (EP^*). Given that $P = {P_n}^{\phi}(EP_m^*)^{1-\phi}$ and $P^* = P_m^*$, we thus have:

$$q' = \frac{P}{EP^*} = {P_n}^{\phi}(EP_m^*)^{-\phi}. \tag{4.18}$$

It can easily be shown that the two real exchange rates are related as follows:

$$q' = q^{\phi}(P_x^*)^{\gamma\phi}(w_x^*)^{-\gamma\phi}(\tfrac{\psi_m}{a_m})^{-\phi}$$

and so we have:

$$\Sigma(q'; p_x^*) = \phi[\Sigma(q; p_x^*) + \gamma],$$

where the symbol Σ denotes an elasticity.

[10] For a discussion of the alternative definitions of the real exchange rates, see for instance Edwards (1989) and Chinn (2006).

4.2.1 Financial openness

In the analysis that follows, financial openness is given by the degree of international capital mobility. To simplify the analysis, we assume in this Subsection that E is fixed (and set arbitrarily equal to 1).

We start by assuming that capital is perfectly mobile internationally. Under this assumption, the domestic rate of interest (r) is tied to the world interest rate (r^*) through the uncovered interest parity condition. With E fixed, we thus have: $r = r^*$, and the domestic economy takes the world interest rate as given. In this case, as shown in numerous papers (see for instance Gregorio and Wolf (1994) or Obstfeld and Rogoff (1996)), the real exchange rate is uniquely determined by Equation (4.6) and thus depends only on the productivity parameters and the world interest rate. Demand conditions have no impact on q. The intuition is straightforward. As P_x and r are given, it follows that the wage rate, w, is determined by Equation (4.4). Furthermore, as shown by Equation (4.5), the wage rate is the only determinant of the price of the nontradable good (P_n). As the real exchange rate is equal to the ratio between the price of the nontradable good and the price of the exportable good, it therefore depends only on P_x^*, a_x, a_n, and r^*. As it appears from Equation (4.6), an increase in the world market price of the exportable commodity leads to an increase (appreciation) of the real exchange rate: $E(q; p_x^*) = \frac{1-\alpha}{\alpha}$. Indeed, an increase in P_x leads to a more than proportional increase in the wage rate and the price of the nontradable good. We can also notice that the lower is α, the stronger is the response of q to a variation in P_x^*.

Let's now assume that capital is not mobile internationally. In this case, r is no longer exogenously given by the world interest rate but is determined endogenously from domestic conditions. The capital stock now becomes exogenous. In this case, q is determined by Equation (4.15) while Equation (4.6) is solved for r. It therefore appears that in response to a change in P_x, q remains unchanged. Indeed, as K_x is now fixed, the production of the tradable good, Y_x, becomes also fixed. The marginal productivity of capital and labor in the tradable sector is therefore fixed which implies that any increase in the price of the exportable good leads to a proportional increase in the domestic rate of interest and the wage rate. In turn, the price of the nontradable good increases in the same proportion, so leaving unchanged the real exchange rate. Alternatively, notice from Equation (4.12) that with Y_x given, the consumer demand of the nontradable good depends only on q. As Y_n is constant, the equilibrium condition for the nontradable good market requires that C_n be constant as well, which imposes that q be invariant.

When we use the alternative definition of the real exchange rate (q'), the reaction of the real exchange rate to a change in the world market price of the primary commodity is given by the following elasticities: (i) $\phi\left(\frac{1-\alpha}{\alpha} + \gamma\right)$

when there is perfect capital mobility and (ii) $\phi\gamma$ in case of zero capital mobility.

The analysis in the Subsection therefore suggests that in more financially open economies, the response of the real exchange rate to a commodity price shock is more pronounced.

4.2.2 Exchange rate regime

The response of the real exchange rate to a commodity price shock when the nominal exchange rate is fixed has been presented in the previous Subsection and we have seen that this response depends on the degree of international capital mobility. When the exchange rate is freely floating, it can be derived from Equations (4.3), (4.6) and (4.15) that the real exchange rate remains unchanged when there is a commodity price shock, whatever the degree of international capital mobility. The explanation is as follows. With E being now endogenous, the model can only be solved provided that r and K_x be exogenous. In this case, to maintain the equality between the marginal cost of capital and its marginal productivity, the price of the exportable good expressed in domestic currency must remain constant. As P_x is unchanged, w and P_n are in turn unchanged and so is q. Accordingly, in the case of an international commodity price shock, the only variable that is affected is the nominal exchange rate: as implied by the law of one price, E appreciates (depreciates) when P_x^* increases (decreases). Our results are therefore consistent with the widespread idea that a flexible exchange rate isolates the domestic economy from international shocks.

When we use the alternative definition of the real exchange rate (q'), the elasticity of q' with respect to p_x^* is equal to $\phi\gamma$, when there is perfect or zero capital mobility.

So, our model suggests that the exchange rate regime only matters when capital is perfectly mobile internationally. In this case, comparing the results of this Subsection with those of the previous Subsection reveals that the real exchange rate is less affected by a commodity price shock when the nominal exchange rate is floating rather than fixed.

4.2.3 Commodity type

In order to address whether the relationship between the real exchange rate and the world market price of the commodity differs according to the type of the commodity that is exported by the small country, we simply consider that what makes a particular commodity different from another one is its technology of production[11]. For instance, it seems reasonable to consider that the produc-

[11] An other way to investigate if the commodity type matters would be to consider that some commodities are available for consumption while others can only be used as intermediate

tion of agricultural commodities is more labor intensive than the production of mineral commodities. We therefore limit our analysis about the impact of the commodity type by examining how the degree of labor intensity (α) affects the elasticity between q (q') and P_x^*. From the analysis above, it turns out that when the exchange rate is flexible or when there is zero capital mobility, the elasticity is independent of α and thus invariant to the type of commodities exported by the country. Conversely, when the exchange rate is fixed and capital is perfectly mobile internationally, the elasticity is a negative function of α. Our model therefore suggests that the impact of international commodity price shocks on the real exchange rate will be lower for countries that are specialized in the export of a commodity whose production is more labor intensive (e.g. agricultural commodities) than for countries which produce capital-intensive commodities (e.g. mineral goods).

4.2.4 Degree of trade openness

In our model, the degree of trade openness can be captured by the parameter (ϕ), which measures the share of the nontradable good, compared to the foreign imported good, in total consumption. It follows from our model that the extent to which the real exchange rate is affected by world commodity price shocks does depend on the country's degree of trade openness only when the real exchange rate is given by q' and when there is zero capital mobility. When those conditions hold, it appears that the higher is the degree of trade openess (the lower is ϕ), the lower is the appreciation (depreciation) of q' in response to a rise (fall) of world commodity prices.

4.2.5 Export diversification

In order to investigate whether the degree of export diversification affects the magnitude of the reaction of the real exchange rate to international commodity price shocks, we extend our model to include the production of a second tradable good. We assume that this new tradable good is a manufacturing good, produced but not consumed domestically. We also assume that the production of the manufacturing good (Y_d) involves two intermediate inputs, the primary commodity (X_d) and an intermediate input produced by the foreign economy (G_d^*):

$$Y_d = a_d {X_d}^{\beta} {G_d^*}^{1-\beta}. \tag{4.19}$$

It is straightforward to show that, in equilibrium, the price of the manufacturing good is determined as follows:

goods

$$P_d = \frac{\psi_d}{a_d} {P_x}^{\beta} {P_f}^{1-\beta}, \tag{4.20}$$

where P_f is the price of the foreign intermediate good in domestic currency with $P_f = EP_f^*$, and $\psi_d = \beta^{-\beta}(1-\beta)^{-(1-\beta)}$.

In this new framework, we now redefine the real exchange rate q as the ratio between the price of nontradable good and a composite price of the two exportable goods, $\tilde{q} = \frac{P_n}{P_t}$, where $P_t = {P_x}^{\epsilon} {P_d}^{(1-\epsilon)}$.

In what follows, the analysis is limited to the case when the exchange rate is fixed and capital is perfectly mobile internationally. Under these assumptions, we obtain through several straightforward substitutions that the elasticity of $\tilde{q}$ with respect to P_x^* is given by:

$$\Sigma(\tilde{q}; p_x^*) = \left(\tfrac{1-\alpha}{\alpha}\right) + (1-\beta)(1-\epsilon).$$

This elasticity is then larger than the corresponding elasticity obtained in the model with one exportable good. So our model does indicate that the degree of export diversification may influence the magnitude of the real exchange rate variation in response to a world commodity price shock. In particular, our model implies that when commodity prices increase (decrease), countries whose exports are largely diversified should register a stronger appreciation (depreciation) of their real exchange rate than countries whose exports are weakly diversified. This results comes from the fact that variation in P_x leads to a less than proportional variation in P_d (see Equation (4.20)).

Notice that the inclusion of a second good does not affect the expression of q' and so, the elasticity $\Sigma(q'; p_x^*)$ is independent of the degree of export diversification.

4.3 Data

In our empirical investigation, we focus on developing and emerging countries that are specialized in the export of a main commodity. We examine whether their real exchange rate is related to the world market price of their main commodity export and whether the relationship is dependent on structural factors. To address those questions, we use annual data covering the period 1980-2007. The selection of the countries and the dataset are discussed in this Section. Data sources are provided in the appendix of this chapter.

4.3.1 Selection of countries and commodities

Our dataset is composed only of countries that are specialized in the export of a leading commodity. Using the results of Bodart, Candelon, and Carpantier (2011), we consider that a country is specialized in the export of a main commodity if that commodity accounts for at least 20 percent of its total exports[12]. From an initial sample of 65 developing and emerging countries that, according to the IMF, are considered as commodity producing countries[13], we found 33 countries that satisfy that criteria. The selected countries are listed in Table 4.1. We indicate in front of each country what is its main commodity export and what is the share of this commodity in the total exports of the country[14]. One can notice that our dataset is mainly composed of African and Latin american countries and include 12 different commodities. More details about the selection of countries are given in the appendix of this chapter.

4.3.2 Real exchange rates and commodity prices

In line with many studies on the determination of the real exchange rate, the exchange rate series is the IMF's real effective exchange rate based on consumer prices. As in Cashin, Cespedes, and Sahay (2004), commodity prices are expressed in real terms, by deflating the US dollar price of each commodity by the IMF's index (of the unit value) of manufactured exports (MUV)[15]. Real exchange rates ($REER$) and real commodity prices ($COMP$) are indices with base January 1995=100.

4.3.3 Structural factors

Each structural factor is given by a three-category variable. We adopted this three-way classification because it permits to assess more accurately the impact of extreme regimes than a two-way classification. In that respect, increasing the number of categories would not be very useful, and it would complicate the econometric estimation.

Exchange rate regime (EXR). As in Broda (2004), we classify the exchange rate regime in three categories: Fixed, Intermediate, and Floating. Our classification is established using mainly the six-way exchange rate regime classification of Ilzetski, Reinhart, and Rogoff (2008), which is an update of

[12] Bodart, Candelon, and Carpantier (2011) finds that a commodity specialization of at least 20 percent is necessary to have cointegration between the real exchange rate and the international price of the leading commodity export

[13] It is the same set of countries as in Cashin, Cespedes, and Sahay (2004), to which we added oil producing countries.

[14] The share is measured as an annual average over the period 1988-2008.

[15] MUV is the unit value index (in US dollars) of manufacturing exports from 20 developed countries with country weights based on the countries' total 1995 exports of manufactures (base 1995=100). The MUV index deflator is taken from the IMF's IFS database.

the Reinhart and Rogoff (2004) classification. Fixed regimes include countries with currency pegs and narrow currency bands. Countries with crawling pegs, wide currency bands or managed floats are included in the intermediate category. The floating category includes countries with freely floating exchange rates. The regimes vary across countries and over time. Table 4.1, columns 5 to 7, report the number of years that each country has spent in every category. Additional details can be found in the appendix of this chapter.

Financial openness (KAO). Using the financial openness index of Chinn and Ito (2006) (hereafter ChI), we classify the countries in 3 categories: Closed, Intermediate, and Open. The 'closed' category includes countries for which the ChI financial openness index is below -1.1. Countries with a financial openness index comprised between -1.1 and 1.0 are included in the 'intermediate' category, while the 'financially open' category includes countries with an index above 1.0. The thresholds used to build our three-way classification are based on the quantile distribution of the Chin-Ito index and follows closely the classification established by Beine, Lodigiani, and Vermeulen (2009). Notice that each country financial openness index is varying over time and Table 4.1, columns 8 to 10, report the number of years that each country has spent in every category.

Trade openness (TRADE). The trade openness variable is also designed as a three-way dummy variable: Closed, Intermediate, and Open. It is based on the GDP ratio of the sum of the total exports and imports of a country. Using the quartiles, we classified a country as closed if its trade ratio was in the first quartile ($< 42.7\%$). The Intermediate category includes the countries whose trade ratio is in the second or third quartiles ($42,7\% < x \leq 88.6\%$), while a country was classified as open to trade if its trade ratio was in the largest quartile ($> 88.6\%$). Table 4.1, columns 11 to 13, report the number of years that each country has spent in every category.

Export diversification (DIV). We distinguish between three categories of export diversification: High, Intermediate, and Low. To determine the categories, we used the data about the export share of the main commodity export (as reported in Table 4.1) and we divided the cross-country distribution in three quantiles. The high diversification category includes countries for which the share of the main commodity in total exports lies between 20% and 31%. Countries with a share lying between 31% and 50% are included in the intermediate category while those with a share higher than 50% are considered as weakly diversified. As the commodity export ratio is computed as an average over the period $1988 - 2008$, the index of export diversification is constant though time and only varies across country units. Ten countries have a High specialization index, twelve countries have an Intermediate specialization index, and eleven countries, most of them being oil producing countries, have a Low specialization index.

Commodity type ($TYPE$). The twelve different commodities that we have in our data set (see Table 4.1) are regrouped in three categories: Energy (oil), Metals (cooper, gold and uranium), and Agriculture (banana, coffee, cocoa, cotton, fish, soya, tea, and tobacco). Sixteen countries are the Energy category, five countries are in the Metal category, and twelve countries are in the Agriculture category.

For ease of convenience, in what follows, the three categories of each structural variable will be referred to as CAT_1, CAT_2, and CAT_3. The three categories of EXR, KAO, $TRADE$, DIV, and $TYPE$, are respectively as follows: CAT_1 = (Fixed, Closed, Closed, High, Energy), CAT_2 = (Intermediate, Intermediate, Intermediate, Intermediate, Metals) and CAT_3 = (Floating, Open, Open, Low, Agriculture).

4.4 Empirical analysis

4.4.1 Preliminary analysis

Before proceeding to the estimation of the relationship between real exchange rates and commodity prices, a preliminary analysis of the data is required. We start testing whether the real exchange rate and commodity prices series are cross-sectional dependent or not. Recent papers show indeed that the consistency of standard panel estimators as well as the size of basic tests (in particular unit root, cointegration) are affected by cross-sectional dependence. We report in Table 4.2 Panel A the outcomes of the Pesaran (2004) which consists in testing for the presence of cross-sectional dependence vs its absence.

It appears that the null hypothesis of cross-sectional independence is strongly rejected by the test. Accordingly, our empirical analysis relies only on panel econometric methods that are robust to cross-sectional dependence. Secondly, we look at the non stationarity of the real exchange rates and commodity prices series. Three tests are performed[16]. According to the results reported in Table 4.2 Panel B, the three tests fail to reject the null of a unit root in the panel data on real exchange rates and commodity prices at usual nominal size , i.e. 5%.

Third, we test whether real exchange rates and commodity prices are cointegrated. To obtain outcomes that are robust to cross-sectional dependence, we use the version of the standard cointegration test of Pedroni (1999) that was developed recently by Fachin (2007). In Table 4.2 Panel C, we report

[16] For a description of the tests, see Bodart, Candelon, and Carpantier (2011)

Table 4.2: Cross-sectional dependence, panel unit root and panel cointegration tests

A- Cross-sectional dependence test				
	test value	p-value	Dependence?	
	45.039	0.00000	Yes	
B- Panel unit root tests				
UR	test value	block rule	crit val (5%)	crit val (10%)
Real exchange rates				
LLC	-12.58	SC	-15.39	-14.27
		MV	-15.78	-14.53
IPS	-2.102	SC	-2.434	-2.409
		MV	-2.50	-2.37
INVN	-0.2043	SC	-2.687	-2.451
		MV	-4.63	-3.29
Commodity prices				
LLC	-9.175	SC	-15.23	-14.66
		MV	-16.27	-13.52
IPS	-1.760	SC	-2.597	-2.461
		MV	-2.58	-2.43
INVN	4.466	SC	-3.18	-3.01
		MV	-6.11	-3.94
C- Panel cointegration tests				
	test value	p-values		
		basic bootstrap	FDB1	FDB2
With common time dummies				
Mean CADF	-3.13	2.00	1.00	2.00
Median CADF	-2.80	22.00	20.00	19.00
Without common time dummies				
Mean CADF	-3.01	4.00	4.00	5.00
Median CADF	-2.84	20.00	17.00	18.00

Notes. The cross-sectional dependence test follows Pesaran (2004). The Pesaran test is based on fixed effects models. The null of the Pesaran test is the absence of cross sectional dependence. The panel unit root tests follow the sub-sampling based approach of Choi and Chue (2007). Real exchange rates and commodity prices are in logarithms. An intercept and a trend have been considered in all the experiments. "SC" and "MV" hold for Stochastic Calibration and Minimum Volatility, respectively, two alternative block selection rules. The minimum and maximum block sizes for the MV rule are 0.4 and 0.6, respectively. An asterisk indicates the rejection of the null of panel unit root (no rejection in our results). The panel cointegration tests are Fachin (2007)'s versions of Pedroni (1999)'s tests. Real exchange rates and commodity prices are in logarithms. An intercept and a trend have been considered in all the tests. Block size selection for the cointegration test is based on 0.1T. "FDB1" and "FDB2" hold for Fast Double Bootstraps of types 1 and 2 (see Davidson and MacKinnon (2006)). P-values are in percent.

both mean and median t-tests, with and without time dummies. The evidence is mixed, as the null hypothesis of no cointegration is rejected by the mean t-tests, but not by the median t-tests. According to the results on individual countries reported in Table 4.1, it seems nevertheless that the presence of a cointegration relationship between real exchange rates and commodity prices cannot be rejected for 17 out of 33 countries. On the basis of this additional evidence, we consider in the rest of our analysis that real exchange rates and commodity prices are panel cointegrated.

The next step in our empirical analysis is to estimate the following cointegration relationship between real exchange rates and commodity prices:

$$REER_{i,t} = a + b.COMP_{i,t} + e_{i,t} \tag{4.21}$$

where $REER_{i,t}$ is the (log) real exchange rate of country i at time t, and $COMP_{i,t}$ is the (log) real price at time t of the leading commodity exported by country i.

Several techniques have been proposed to estimate such a long run relationship, the most popular being DOLS and FMOLS. Nevertheless Bai, Kao, and Ng (2009) (BKN hereafter) prove that in presence of cross-sectional dependence generated by unobserved global stochastic trends traditional estimators are biased. They introduce common factors to control for cross unit dependence, and thus propose an iterative procedure to extract the common factor and to estimate the equation simultaneously. The result are reported in Table 3 column 1. It turns out that the price of the dominant commodity has a long-run impact on the real exchange rate with an elasticity of 12.5% which is in line with the Bodart, Candelon, and Carpantier (2011)'s analysis, the only one so far to implement BKN approach in this context.

4.4.2 Estimates of the impact of structural factors

Once this preliminary analysis achieved, let us now turn to the paper's core, i.e. how the selected structural factors might affect the long-run commodity price elasticity of the real exchange. To this aim, Equation (4.21) is augmented by an interaction term between the commodity price variable and each structural factor, given by a three-category variable (see Section 4.3). Hence, we explore their impact by interacting the commodity price variable with two dummy variables, a dummy D_2 (that takes the value 1 when observations are in the 2nd category (CAT_2) and zero otherwise) and a dummy D_3 (that takes the value 1 when observations are in the 3rd category (CAT_3) and zero otherwise). Let us note that we did not include a D_1 dummy to avoid the linear dependence between the interaction terms and the variable $COMP_{i,t}$. The category defined by D_1 will thus be considered as the normalization. This gives the following relationship:

$$REER_{i,t} = a + b.COMP_{i,t} + c_1(D_{j,2,t} * COMP_{i,t}) + c_2(D_{j,3,t} * COMP_{i,t}) + e_{i,t} \quad (4.22)$$

The estimation of Equation (4.22) thus amounts to estimate the long-run relationship between real exchange rates and commodity prices, conditional on each factor. The impact of each factor on the long-run commodity price elasticity of the real exchange rate is determined by testing the hypothesis that each c_j $(j = 1, 2)$ is significantly different from zero ($H0$: $cj = 0$).

The estimation outcomes of Equation (4.22) are reported in Table 4.3 columns (2) to (6). Before analyzing the impact of the structural factors on the relationship between real effective exchange rate and the price of the main primary commodity, let us make two remarks. First, it is noticeable that the baseline coefficient b only varies marginally from one specification to the other, lying between 0.101 and 0.162, highlighting hence the robustness of the elasticity. Second, intermediate category variables always appear as non significant. It confirms hence that splitting the structural factors in 3 regimes is sufficient to apprehend their potential roles, and that only extreme categories (flexible exchange rate, trade openness, weakly diversification) lead to a significant deviation of the elasticity from the one obtained with the baseline model.

As concerns the effects of the structural factors, outcomes from the regressions confirm our theoretical predictions, i.e. a flexible exchange rate, the trade openness and a high degree of diversification[17] decreases the elasticity between the real effective exchange rate and the price of the main exported raw commodity. Regarding the impact of the exchange rate regime, our results question the empirical evidences of Chen and Rogoff (2003) and Broda (2004) who, conversely, find that the impact of world commodity prices on the real exchange rate is weaker when the exchange rate is fixed. We also find that when the most exported raw commodity is an agricultural good, the elasticity becomes higher, corroborating our finding that real exchange rate and commodity prices are strongly linked for labor intensive activity sectors. Interestingly, column (3) does not show any effect of financial openness. This result, which does not match the predictions of theoretical models, may be due to the difficulty to measure properly this structural variable. However, to our knowledge, the Chinn and Ito (2006) index remains the most accurate and used measure of financial openness.

[17]Table 4.3 row 9 reports that a weak diversification increases significantly the elasticity.

Table 4.3: Estimation results

coef.	(1)	(2)	(3)	(4)	(5)	(6)
b	0.125***	0.137***	0.129***	0.139***	0.101**	0.162***
	(0.022)	(0.022)	(0.023)	(0.023)	(0.041)	(0.027)
$c_{EXR,2}$		-0.008				
		(0.007)				
$c_{EXR,3}$		-0.044***				
		(0.009)				
$c_{KAO,2}$			-0.005			
			(0.005)			
$c_{KAO,3}$			0.001			
			(0.010)			
$c_{TRADE,2}$				-0.008		
				(0.006)		
$c_{TRADE,3}$				-0.019**		
				(0.009)		
$c_{DIV,2}$					-0.006	
					(0.054)	
$c_{DIV,3}$					0.132**	
					(0.055)	
$c_{TYPE,2}$						0.028
						(0.057)
$c_{TYPE,3}$						-0.100*
						(0.056)

Notes. Column (1) reports the results of Equation 4.21 where real exchange rates are regressed on commodity prices, and where b is the elasticity. Columns (2)-(6) report the estimates of Equation 4.22 for the different factors added in the specification. The factors are respectively the exchange rate regime (EXR), the financial openness (KAO), the trade openness (TRADE), the export diversification (DIV) and the commodity type (TYPE). The subscripts 2 and 3 in the label of the factors refer to the three-way classification, where 2 stands for intermediate, intermediate, intermediate, intermediate and metals, respectively for EXR, KAO, TRADE, DIV and TYPE, and where 3 stands for floating, open, open, low and agriculture, respectively for EXR, KAO, TRADE, DIV and TYPE. Standard errors are reported in parentheses. *, ** and *** stand for significance at 10%, 5% and 1%, respectively.

4.4.3 Robustness check

To evaluate the robustness of our findings, three investigations are performed. In a first step, the specification of the dummy variables representing the trade openness and the export diversification might be criticized (the classification of the exchange rate regime, the financial openness and the type of commodities is not subject to debate). We thus scrutinized in a first step the robustness of our results to alternative definition of $TRADE$ and DIV. Instead of being classified using quartile, the new trade dummy $TRADE^n$ is now built from "quintiles". It results that to be considered as open a country should have an openness ratio higher than 94%, whereas it is considered as closed if it lies below 38%. Similarly, the new DIV^n dummy is built upon quartiles. Results of the estimations are reported in Table 4.4. It appears that the results are qualitatively comparable to those previously obtained (Table 4.3 columns (4) and (5)) supporting hence the robustness of our findings to the definition of the structural factors.

Table 4.4: Estimation results with modified trade openness and export diversification dummies

	b	$c_{i,2}$	$c_{i,3}$
$i = TRADE^n$	0.161*** (0.023)	-0.013** (0.006)	-0.023** (0.009)
$i = DIV^n$	0.125*** (0.045)	0.004 (0.054)	0.134** (0.060)

Notes. Standard errors in parentheses. *, ** and *** stand for significance at 10%, 5% and 1%, respectively.

In a second step, countries are split into two groups indicating their geographical areas (Sub-Saharan Africa and Latin America). Estimations (1) to (6) are repeated for each of these groups and are reported in Tables 4.5 and 4.6. It turns out that the obtained coefficients are qualitatively identical to those reported in Table 4.3 corroborating our conclusions and also stressing that they also hold at the regional level. A small discrepancy is nevertheless observed for coefficients associated with the type of commodity: they are either non significant for Sub-Saharan either positive for Latin America. It is possible to explain this feature by the low diversification in the type of commodities. For example in the Sub-Saharan panel, half of countries have an agricultural commodity as their main commodity export.

Finally, in a third step, we control for devaluations i.e. sharp nominal exchange rate adjustment. In particular, many African countries have been affected by the Franc CFA devaluation in January 1994. Table 4.7 reports the results of the estimations, when taking into account these devaluations. It is noticeable

Table 4.5: Estimation results for Sub-Saharan African countries

coef.	(1)	(2)	(3)	(4)	(5)	(6)
b	0.210***	0.151***	0.163***	0.155***	0.039	0.157***
	(0.036)	(0.027)	(0.029)	(0.030)	(0.046)	(0.037)
$c_{EXR,2}$		0.003				
		(0.010)				
$c_{EXR,3}$		-0.068***				
		(0.012)				
$c_{KAO,2}$			0.010*			
			(0.006)			
$c_{KAO,3}$			0.040***			
			(0.012)			
$c_{TRADE,2}$				0.004		
				(0.008)		
$c_{TRADE,3}$				-0.030**		
				(0.012)		
$c_{DIV,2}$					0.030	
					(0.079)	
$c_{DIV,3}$					0.221***	
					(0.061)	
$c_{TYPE,2}$						-0.046
						(0.082)
$c_{TYPE,3}$						0.016
						(0.064)

Notes. Countries included are Burundi, Cameroon, Cote d'Ivoire, Ethiopia, Gabon, Ghana, Kenya, Malawi, Mali, Mauritania, Niger, Nigeria, Sudan, Tanzania, Uganda and Zambia. Column (1) reports the results of Equation 4.21 where real exchange rates are regressed on commodity prices, and where b is the elasticity. Columns (2)-(6) report the estimates of Equation 4.22 for the different factors added in the specification. The factors are respectively the exchange rate regime (EXR), the financial openness (KAO), the trade openness (TRADE), the export diversification (DIV) and the commodity type (TYPE). The subscripts 2 and 3 in the label of the factors refer to the three-way classification, where 2 stands for intermediate, intermediate, intermediate, intermediate and metals, respectively for EXR, KAO, TRADE, DIV and TYPE, and where 3 stands for floating, open, open, low and agriculture, respectively for EXR, KAO, TRADE, DIV and TYPE. Standard errors are reported in parentheses. *, ** and *** stand for significance at 10%, 5% and 1%, respectively.

Table 4.6: Estimation results for Latin American countries

coef.	(1)	(2)	(3)	(4)	(5)	(6)
b	0.089***	0.120***	0.102***	0.224***	-	0.122***
	(0.028)	(0.031)	(0.034)	(0.031)	-	(0.034)
$c_{EXR,2}$		-0.006				
		(0.012)				
$c_{EXR,3}$		-0.029**				
		(0.013)				
$c_{KAO,2}$			0.006			
			(0.008)			
$c_{KAO,3}$			0.013			
			(0.014)			
$c_{TRADE,2}$				-0.063***		
				(0.008)		
$c_{TRADE,3}$				-0.058***		
				(0.012)		
$c_{DIV,2}$					-	
					-	
$c_{DIV,3}$					-	
					-	
$c_{TYPE,2}$						0.211**
						(0.088)
$c_{TYPE,3}$						0.145
						(0.104)

Notes. Countries included are Chile, Colombia, Dominica, Ecuador, Honduras, Paraguay and Venezuela. Column (1) reports the results of Equation 4.21 where real exchange rates are regressed on commodity prices, and where b is the elasticity. Columns (2)-(6) report the estimates of Equation 4.22 for the different factors added in the specification. The factors are respectively the exchange rate regime (EXR), the financial openness (KAO), the trade openness (TRADE), the export diversification (DIV) and the commodity type (TYPE). The subscripts 2 and 3 in the label of the factors refer to the three-way classification, where 2 stands for intermediate, intermediate, intermediate, intermediate and metals, respectively for EXR, KAO, TRADE, DIV and TYPE, and where 3 stands for floating, open, open, low and agriculture, respectively for EXR, KAO, TRADE, DIV and TYPE. Standard errors are reported in parentheses. *, ** and *** stand for significance at 10%, 5% and 1%, respectively.

that the estimators are very closed to those obtained in Table 4.3, supporting again the robustness of our conclusions.

Table 4.7: Estimation results conditional to devaluations

coef.	(1)	(2)	(3)	(4)	(5)	(6)
b	0.143***	0.145***	0.141***	0.149***	0.114***	0.167***
	(0.021)	(0.021)	(0.022)	(0.023)	(0.040)	(0.026)
$c_{EXR,2}$		-0.005				
		(0.007)				
$c_{EXR,3}$		-0.037***				
		(0.009)				
$c_{KAO,2}$			0.000			
			(0.005)			
$c_{KAO,3}$			0.008			
			(0.010)			
$c_{TRADE,2}$				-0.003		
				(0.006)		
$c_{TRADE,3}$				-0.012		
				(0.009)		
$c_{DIV,2}$					-0.015	
					(0.053)	
$c_{DIV,3}$					0.127**	
					(0.054)	
$c_{TYPE,2}$						0.041
						(0.055)
$c_{TYPE,3}$						-0.097*
						(0.055)

Notes. Column (1) reports the results of Equation 4.21 where real exchange rates are regressed on commodity prices, and where b is the elasticity. Columns (2)-(6) report the estimates of Equation 4.22 for the different factors added in the specification. The factors are respectively the exchange rate regime (EXR), the financial openness (KAO), the trade openness (TRADE), the export diversification (DIV) and the commodity type (TYPE). The subscripts 2 and 3 in the label of the factors refer to the three-way classification, where 2 stands for intermediate, intermediate, intermediate, intermediate and metals, respectively for EXR, KAO, TRADE, DIV and TYPE, and where 3 stands for floating, open, open, low and agriculture, respectively for EXR, KAO, TRADE, DIV and TYPE. Standard errors are reported in parentheses. *, ** and *** stand for significance at 10%, 5% and 1%, respectively.

4.5 Conclusions

As it is demonstrated by the literature on the "Dutch" disease, commodity prices are potentially an important determinant of the real exchange rate of countries that produce and export primary commodities. For that reason, the dependence of real exchange rates on commodity prices (or, more generally, on terms of trade) has been the focus of numerous empirical studies. In most of these studies, the empirical evidence consists only of econometric estimates of the real exchange rate response to commodity price shocks. What determines the magnitude of the real exchange rate reaction is however not investigated. This issue is the main subject of this paper. We first showed with a simple theoretical framework that several structural or institutional characteristics may contribute to shape the relationship that exists in the long-run between a country's real exchange rate and the world price of the main commodity that it exports. We emphasized the role of five structural factors: the exchange rate regime, the degree of trade openness, the degree of export diversification, the type of the commodity export, and the degree of financial openness. We then subjected the role of these factors to econometric scrutiny by conducting an empirical investigation involving 33 developing countries specialized in the export of a main primary commodity.

Using panel cointegration methods robust to cross-sectional dependence, we found that four of the five factors, namely the exchange rate regime, the degree of trade openness, the degree of export diversification and the type of the commodity exported, have a significant impact on the long-run commodity price elasticity of the real exchange rate. More precisely, we report that the elasticity is reduced when a country operates a floating exchange rate (rather than a fixed exchange rate). It is also smaller when the country is strongly opened to external trade (rather than closed) and when its exports are highly diversified (rather than specialized on a few products). We also find that the elasticity is smaller when the primary commodity is an agricultural good, but is higher when it is oil. Conversely, it appears from our analysis that the degree of financial openness does not influence significantly the extent to which real exchange rates and world commodity prices are related in the long-run. We finally note that these results also hold at the regional level and are not affected by sharp devaluations (such as the franc CFA devaluation of January 1994).

Our results have several policy implications. They suggest in particular that countries which are concerned by the adverse impact that world commodity price fluctuations may have on their competitiveness, in other words countries that wish to reduce "Dutch disease" effects, should adopt a floating exchange rate instead of a fixed exchange rate. They also suggest that developing countries whose exports are concentrated on a few primary commodities can reduce the dependance of their real exchange rate on world commodity prices by having more diversified exports. According to our results, diversification towards agricultural products would isolate further the real exchange rate from world commodity price shocks. Finally, countries can also protect their competitiveness from fluctuations in world commodity prices by being more opened to external trade.

Appendix: description of the data

To make the selection of the countries included in our dataset, we recovered from the UN Comtrade database the annual US$ export value of 42 commodities for 65 countries over 21 years (1988-2008). The set of countries corresponds to the sample of developing and emerging countries selected by Cashin, Cespedes, and Sahay (2004) on the basis of the International Monetary Fund classification of developing countries according to the composition of their export earnings, to which we added oil producing countries. For each country, we expressed the 1988-2008 annual average of the US$ export receipt of each individual commodity as a share of the total US$ export receipts of the country. We then selected the countries where it appeared that at least one commodity had a share in total exports of at least 20 percent. From the resulting list of countries, a few of them were eliminated because we detected that the commodity export share was very volatile over the years, most probably because of data reporting deficiencies. Precisely, a country was excluded from our analysis when the export share of dominant export was less than 2% during 5 years or more. For example, Mozambique was removed because the export share of its aluminium exports was more than 50% after 2001 but about zero before 2000. A few other countries were excluded from our analysis because of the unavailability of data for all the variables of our model.

Data on the real exchange rate are the International Monetary Fund real effective exchange rates based on consumer prices[18]. The data are extracted from the IMF's Information Notice System (INS) database.

Commodity prices are taken from the International Financial Statistics (IFS) database of the IMF. Two series, Tobacco and Gold, were not available in IFS and were taken from Datastream (with respective codes: USI76M.ZA and GOLDBLN). Details about each series are provided in Table 4.2.

Data on exchange rate regimes are taken from two databases. The first database, *IRR*, is provided by Ilzetski, Reinhart, and Rogoff (2008) and is an update of the Reinhart and Rogoff (2004). The *IRR* exchange rate classification is based on market-determined parallel exchange rates. It provides monthly and annual classification data over the period 1940-2007. Classification by 6 and 15 categories are available. The second database,

[18]See Desruelle and Zanello (1997) for details regarding the construction of the real effective exchange rates.

LYS, was established by Levy-Yeyati and Sturzenegger (2005) on the basis of data of interventions (volatility of reserves and exchange rates). It covers 183 countries for the period 1974-2004. Classification by 3 and 5 categories are available. Both databases have advantages and weaknesses. One advantage of the LYS database is to provide a 3-category classification, but it has no data for 2005 and later. On the contrary, the IRR database covers the period 1980-2007, but it has only 6- and 15-category classifications. Our paper uses mainly the IRR 6-regime classification to benefit from the extended coverage 1980-2007. It contains however missing values. When possible, we solved this problem by replacing the missing value by the value found in the LYS 3-category classification. When no data were available in the LYS database, we replaced the missing value by the value of the previous year. We converted the IRR6-regime classification into three exchange rate categories as follows: our Fixed Regime category corresponds to IRR category 1; our Intermediate Regime category corresponds to categories 2 and 3 of IRR; and our Flexible Regime category includes categories 4, 5 and 6 of IRR.

The financial openness dummy variable used in our analysis is based upon the financial openness index of Chinn and Ito (2006). The Chinn-Ito index measures the degree of openness of a country's capital account. It is based on the binary dummy variables that codify the tabulation of restrictions on cross-border financial transactions reported in the International Monetary Fund's *Annual Report on Exchange Arrangements and Exchange Restrictions*. This index is thus a de jure measure of financial openness. The index is not smoothly distributed. Values range between -1.9 and 2.6, with about one third of the observations having a value of -1.13. Three values were missing in the Chinn-Ito database: Dominica in 1980 and 1981 and Sudan in 2006. We filled the gaps by setting the 1980 and 1981 Dominica indices equal to the 1982 Dominica index, and by setting the 2006 Sudan index equal to the 2005 Sudan index (which is the same as the 2007 index).

The GDP, export and import data used to construct the trade openness dummy variable are all taken from the World Bank. All the data are in current US$.

Table 4.8: Description of the commodity price series

Commodity	Source	Description
Bananas	IFS	Central American and Ecuador, first class quality tropical pack, Chiquita, Dole and Del Monte, U.S. importer's price FOB U.S. ports (Sopisco News, Guayaquil). \$/Mt
Cocoa beans	IFS	International Cocoa Organization cash price. Average of the three nearest active futures trading months in the New York Cocoa Exchange at noon and the London Terminal market at closing time, CIF U.S. and European ports (The Financial Times, London). \$/Mt
Coffee	IFS	International Coffee Organization, Other Mild Arabicas New York cash price. Average of El Salvador central standard, Guatemala prime washed and Mexico prime washed, prompt shipment, ex-dock New York. Cts/lb
Copper	IFS	London Metal Exchange, grade A cathodes, spot price, CIF European ports (Wall Street Journal, New York and Metals Week, New York). Prior to July 1986, higher grade, wirebars, or cathodes. \$/Mt
Cotton	IFS	Cotton, Liverpool Index A, cif Liverpool US cent/pound
Fish	IFS	Fresh Norwegian Salmon, farm bred, export price (NorStat). US\$/kg
Gold	DS	Gold Bullion LBM US\$/Troy Ounce
Oil	IFS	U.S., West Texas Intermediate 40o API, spot price, FOB Midland Texas (New York Mercantile Exchange, New York). (In 1983-1984 Platt's Oilgram Price Report, New York). \$/bbl
Soya	IFS	Soybean U.S., cif Rotterdam US\$ per Metric Ton
Tea	IFS	Mombasa auction price for best PF1, Kenyan Tea. Replaces London auction price beginning July 1998. Cts/Kg
Tobacco	DS	Tobacco, US (all markets), mid month curn
Uranium	IFS	Metal Bulletin Nuexco Exchange Uranium (U3O8 restricted) price. \$/lb

Notes."IFS" refers to International Financial Statistics from the IMF. "DS" refers to Datastream. "LME" refers to London Metal Exchange. "bbl" refers to barrel (42 US Gallons). "API" refers to American Petroleum Institute.

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