Real Exchanges Rates in Commodity Producing Countries: A Reappraisal

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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

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1 Introduction

Many developing countries, gifted in natural resources, heavily depend on international commodity prices. Tobacco, for example, represents about 60% of Malawi commodity exports, oil 90% of Nigeria exports. Whether natural resources are a luck or a curse for these countries is still an open debate. Some countries have admirably well taken profits of their endowments, while others suffered countless economic difficulties. Whatever the records, it is generally recognized that commodity prices can be a source of macroeconomic instability in developing countries.

The natural resource curse literature, surveyed recently by Frankel (2010a), identifies several channels through which commodity prices can affect the economic situation of commodity producing countries. One important channel is the real exchange rate. 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 estimate of the relationship between commodity prices and real exchange rates for a large group of developing countries. Recent estimates of this relationship are provided by Chen and Rogoff (2003) and Cashin et al. (2004). The former provides results for 3 developed countries (Canada, Australia and New-Zealand) and the latter for several developing countries. Both use time-series cointegration techniques and find that commodity prices have a significant positive long-run impact on the real exchange rate of many countries.

In their analysis, Chen and Rogoff (2003) and Cashin et al. (2004) look at the impact that commodity prices can have on real exchange rates by using for each country a commodity price index constructed as a weighted average of the price of the different commodities that are produced and exported by the country. Accordingly, they do not pay a particular attention to countries that are specialized in the export of one leading commodity, as for instance Burundi whose coffee exports count for about 50% of its total exports of merchandises or Mali where gold counts for more than 54% of its total exports. For several reasons, we do believe however that it is relevant to investigate the particular case of countries highly specialized in the export of one leading commodity and determine what is the impact of the price of the leading commodity that they export on their real exchange rate. First, a high specialization implies that commodity price fluctuations lead to large variations in the external trade balance. That is the reason why Frankel and Saiki (2002) made the proposal that countries specialized in the export of a particular commodity should peg their currency to the price of that commodity. This mechanism would permit an automatic accommodation of terms of trade shocks. Second, the use of an aggregate price export index instead of a single price implies that variations in the price index reflect not only changes in the individual commodity prices but also changes in the respective weights of the commodities. In addition, when an aggregate price index is used, the relationship that is estimated between the real exchange rate and the commodity price index depends on the correlation among individual prices over the sample period. Accordingly, the true relationship between real echange rates an commodity prices is blurred. For instance, if the aggregate price index is composed of two prices that are perfectly negatively correlated, it is impossible to detect any relationship between the real exchange rate and the aggregate price index. One way to avoid these problems is to use a single commodity price instead of an aggregate price index. For all these reasons, contrary to previous research, our analysis will focus on developing countries that are specialized in the export of one leading commodity.

Several previous papers have been interested in the determination of the real exchange rate of countries with a strong export specialization. This is for instance the case of Edwards (1986), who studies the determination of the real exchange rate in Colombia, a world leading exporter of coffee; of Habib and Kalamova (2007), who concentrate on oil exporting countries; of Sjaastad and Scacciavillani (1996), whose analysis focuses on gold exporting countries; or of Frankel (2007), who investigate whether the price of gold affects the real exchange rate of the South-African rand. Our research is the first, to our knowledge, to examine this issue for a large sample of developing countries.

Our analysis of the relationship between real exchange rates and commodity prices relies non-stationary panel cointegration methods of second generation, i.e. 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.

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 et al. (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 2 presents briefly the empirical literature devoted to the relation between commodity prices and real exchange rates. Section 3 presents the model. Section 4 describes the data. Section 5 presents the methods. Section 6 discusses the empirical results. Some robustness checks are performed in section 7. Section 8 concludes.

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. Though 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. Recent empirical evidence provide support to this view. Chen and Rogoff (2003) show for instance that commodity prices have a strong and positive influence on the real effective exchange rate of Canada, Australia and New-Zealand. Similarly, Cashin et al. (2004) search for a long-run equilibrium relationship between real commodity prices and the real effective exchange rate of 58 commodity-exporting countries and found evidence of such a relationship for about one-third of their sample of countries. Coudert et al. (2008) report similar evidence for a large group of commodity exporting countries, including oil producers.

Following the research initiated by Chen and Rogoff (2003) and Cashin et al. (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 et al. (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.[1] In the most recent papers, for instance Coudert et al. (2008) or Chen and Chen (2007), panel cointegration techniques are used, whereas the results of Chen and Rogoff (2003) and Cashin et al. (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 Model

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

$$REER_{i,t} = \alpha_i + \beta COM_{i,t} + \epsilon_{i,t} \tag{1}$$

where $REER_{i,t}$ is the real effective exchange rate (in logarithm) of country i, $COM_{i,t}$ the price of 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 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 (1) will be estimated using panel cointegration methods with monthly data covering the period 1988-2008.

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.

4.1 Selection of the Relevant Commodity-Country Pairs

In order to detect the countries and commodities 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 et al. (2004), with the addition of oil. The list of commodities is reported in Table 1. The countries included in our sample are the developing and emerging countries selected by Cashin et al. (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 2.

As explained in the previous sections, what is new in our approach is that we relate the real exchange rate of an individual 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%.

2. Using these commodity export ratios, we then determine for each commodity which country is the most highly specialized in the export of that particular commodity. We then use the country that is so identified to obtain one particular country-commodity pair. For example, this procedure shows that Mali has the highest ratio of cotton exports among the countries in our sample. We applied this procedure to the 42 commodities of our sample and we thus obtained a series of 42 country-commodity pairs.

3. We then eliminate from the series the pairs where the commodity share is less than 20% of total commodity exports. This threshold is set arbitrarily but other thresholds will be used as robustness checks.

4. It appeared that, for some countries, the export share of the dominant commodity was very volatile over time. We then decided to eliminate a country-commodity pair when the share of its dominant export was less than 2% during 5 years or more. We removed for instance Mozambique, because the share of its aluminium exports in total exports was more than 50% after 2001 but about zero before 2000.

In Table 3, we report for several 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 country-commodity pairs are kept for our empirical analysis. They are listed in Table 4. The pair Cotton-Benin is not reported due to the unavailability of real exchange rate data over the period 1980-2008. In the last section, we do perform robustness checks using alternative samples.

An important feature about the construction of our sample of countries is that only one country is associated to each commodity. It follows 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 sample. By proceeding like this, we avoid to give too much weight to some commodities, like oil, and so make sure that our results are not influenced by a few number of commodities.

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 5. All the series are available at the monthly frequency over the period 1980-2008.

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).[2] 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 et al. (2004)). The MUV index is represented in Figure 3. Real commodity prices are then normalized with base period January 1995=100. Normalized real commodity prices are represented in Figure 1.

We can see in Figure 1 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.

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.[3] The data come from the IMF's Information Notice System (INS).[4]

We represent the CPI-based real effective exchange rates (normalized to January 1995=100) in Figure 2 for the period 1980 to 2008. 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 4 and 5. One can observe that some series seem to be clearly correlated, while others are not. Formal statistical tests are performed later.

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 6. We see that there are big differences in the standard deviation across commodity prices. For example, the standard deviation of tobacco prices is more than nearly ten times lower than the one of uranium. We may also observe differences in the variability of real exchange rate across countries, the standard deviation of the real exchange rate of Nigeria being more than ten times larger than the one of Dominica. We find unit roots for most of the series (excepted banana, soya and tea prices) using Dickey-Fuller tests (with constant and trend, or just a constant if the trend is not significant).

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 et al. (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 et al. (2004), we considered that these methods were not relevant for our application because the number of observations per country was too small to guarantee that the unit-root based tests reach a good power. We therefore use panel methods to perform our econometric analysis.

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 7. We therefore use second generations panel techniques, which allow for cross-sectional dependence. We briefly describe these methods in the following subsections.[5]

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 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 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 et al. (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.

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 et al. (2008).

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 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 (2000) 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), whic is robust to cross-sectional dependence. They consider the following framework:

$$
\begin{aligned}
REER_{i,t} &= \beta COM_{i,t} + e_{i,t} && (2)\\
COM_{i,t} &= COM_{i,t-1} + \epsilon_{i,t} && (3)\\
e_{i,t} &= \lambda_i' F_t + u_{i,t} && (4)\\
F_t &= F_{t-1} + \eta_{i,t}, && (5)
\end{aligned}
$$

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.

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.

6.1 Stationarity analysis

Output of the stationarity tests of commodity prices is reported in Table 8. 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 with stochastic calibration rule, the null of non-stationarity is rejected. 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, 2004).

Panel unit root tests for the real exchange rates are reported in Table 9. 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.

6.2 Cointegration analysis

Results of Fachin (2006)'s panel cointegration tests are reported in Table 10. We report both mean and median t-tests, based on Pedroni (1999), with and 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 suppoted by our data. Treating each country individually, Cashin et al. (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.

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 11.

Whatever technique is used (DOLS, FMOLS, BKN), it tunrs out that the cointegration coefficient is significant and positive, as expected. The coefficient estimates obtained with DOLS and FMOLS are respectively 0.326 and 0.317, 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 et al. (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 0.149, 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.

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 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%. Results obtained with these new samples are reported in Table 13. They show that cointegration holds from a threshold of about 20%. Indeed, when the threshold is set at 10% and 15%, we cannot reject the null of no cointegration between the commodity price of the leading commodity and the real exchange rate (see Table 12.) In addition, the results in Table 13, which are illustrated in Figure 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.

We found in Table 11 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 14 and confirm that neglecting cross-sectional dependence leads to a higher elasticity estimation.

We also assessed the robustness of our results to the presence of structural breaks 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 15). Outliers are removed by simply setting the monthly change equal to zero. We then reiterate our tests with the modified dataset. As reported in Tables 16 and 17, the results are similar to those obtained with the original dataset. To further check the potential impact of the breaks, 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. Our results remain similar as reported in Tables 16 and 17.

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

Notes

1. See also Chen and Chen (2007) for the relation between oil and importing-countries real

exchange rates.

2. 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.

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

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

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

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Table 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 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	

Table 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%
Alu	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 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	Alu	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 5: Description of the commodity price series

Commodity	Source	Description
Alu	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.

Table 6: Basic descriptive statistics

Series	min	mean	max	std.dev	ADF(p)	p
Nigeria	3.4853	4.4486	6.0982	0.68673	-1.656	0
Oil	4.2297	5.1401	6.3455	0.4649	-2.288	1
Malawi	4.4747	5.1098	5.5274	0.29195	-2.002	1
Tobacco	4.5073	4.8291	5.1419	0.13416	-2.376	1
Zambia	3.241	4.6771	5.3882	0.31473	-2.573	1
Copper	3.9584	4.4385	5.5381	0.35539	-1.863	8
Mali	4.5016	5.0082	5.7875	0.41613	-1.748	7
Gold	4.3503	4.7661	5.4543	0.24692	-1.875	1
Burundi	4.1084	4.6251	5.2196	0.28362	-1.088	0
Coffee	3.3921	4.272	5.2334	0.3995	-2.059	5
Niger	4.2787	4.9547	5.7179	0.36193	-1.758	2
Uranium	4.2447	5.1067	7.0311	0.60505	-1.189	10
Cote d'Ivôire	4.3603	4.8014	5.0642	0.14849	-2.656	0
Cocoa	4.204	4.8064	5.6788	0.35895	-1.862	2
Paraguay	4.270	4.6535	5.2845	0.25216	-2.312	1
Soya	4.5049	4.8607	5.2845	0.25216	-3.847**	1
Mauritania	4.0602	4.6674	5.282	0.36798	-0.9941	0
Fish	4.0098	4.7319	5.4332	0.36975	-1.634	6
Dominica	4.3684	4.5486	4.7373	0.072899	-2.634	0
Banana	4.3005	4.99	5.5221	0.22919	-3.050*	9
Kenya	4.0221	4.5563	5.1507	0.19515	-0.3104	4
Tea	4.4078	4.9566	5.9459	0.24846	-3.194*	2

Notes. Series are in logarithm, normanlized and, as regards commodity prices, deflated by manufacture unit value index (MUV). Critical values for the Augmented Dickey-Fuller tests (with constant) are -2.87 for 5% and -3.45 for 1%.

Table 7: Pesaran test of cross-sectional dependence for different panels of countries

# countries	Pesaran Test	p-value	Dependence?	AACR
5	21.66	0.00	Yes	0.406
10	36.929	0.00	Yes	0.420
15	36.929	0.00	Yes	0.445

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, 10 and 15 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.

Table 8: 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	-7.718	SC	-8.940	68	-8.616	66
		MV	-8.665	32	-8.313	17
IPS	-2.638	SC	-2.518*	48	-2.351*	32
		MV	-2.470	24	-2.351	32
INVN	-0.0360	SC	-2.078	117	-2.109	143
		MV	-2.578	30	-2.349	29

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.

Table 9: 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	-6.479	SC	-9.241	68	-7.785	66
		MV	-11.428	16	-10.157	17
IPS	-2.269	SC	-2.726	48	-2.556	32
		MV	-2.848	20	-2.676	20
INVN	-0.5777	SC	-3.055	117	-3.190	143
		MV	- 4.423	15	-4.062	15

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.

Table 10: Fachin (2006) Panel Cointegration Tests

	Test value	p-values		
		basic bootstrap	FDB1	FDB2
With common time dummies				
Mean CADF	-3.25	2.00	3.00	4.00
Median CADF	-3.28	0.00	1.00	1.00
Without common time dummies				
Mean CADF	-3.21	0.00	0.00	0.00
Median CADF	-3.22	1.00	0.00	1.00

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 of types 1 and 2 (see Fachin (2006)). P-values are in percent.

Table 11: Cointegration esimates by DOLS, FMOLS and BKN methods

Test	β	Standard errors
Panel DOLS	0.326*	0.0022
Panel FMOLS	0.317*	0.0021
BKN	0.149*	0.0092

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 et al. (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.

Table 12: Fachin(2006)'s cointegration tests for different samples

	Test	Stat	p-Btst	FDB1	FDB2	Coint.?
14 countries (> 10%)						
With common time dummies	Mean CADF	-2.84	18	22	21	No
	Median CADF	-2.7	34	42	43	No
Without common time dummies	Mean CADF	-3.04	2	2	3	yes
	Median CADF	-3.09	1	1	1	yes
12 countries (> 15%)						
With common time dummies	Mean CADF	-2.95	7	10	10	yes
	Median CADF	-2.65	37	37	41	No
Without common time dummies	Mean CADF	-3.13	1	1	2	yes
	Median CADF	-3.22	0	0	0	yes
11 countries (> 20%)						
With common time dummies	Mean CADF	-3.29	0	0	-1	yes
	Median CADF	-3.23	2	3	4	yes
Without common time dummies	Mean CADF	-3.25	1	0	1	yes
	Median CADF	-3.53	0	4	0	yes
9 countries (> 30%)						
With common time dummies	Mean CADF	-3.33	3	7	5	yes
	Median CADF	-3.24	3	6	5	yes
Without common time dummies	Mean CADF	-3.42	0	0	0	yes
	Median CADF	-3.54	0	10	0	yes
6 countries (> 40%)						
With common time dummies	Mean CADF	-3.32	4	5	6	yes
	Median CADF	-3.16	4	5	5	yes
Without common time dummies	Mean CADF	-3.7	0	3	0	yes
	Median CADF	-3.66	0	1	0	yes
5 countries (> 50%)						
With common time dummies	Mean CADF	-3.48	2	2	3	yes
	Median CADF	-3.17	8	9	11	yes
Without common time dummies	Mean CADF	-3.74	0	0	0	yes
	Median CADF	-3.77	0	0	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 13: Cointegration tests and estimates for different samples

Weight of the main cdty	# countries	Cointegration tests	Coefficients	SE
More than 50%	(5 countries)	C	0.284	0.0131
More than 40%	(6 countries)	C	0.242	0.0101
More than 30%	(9 countries)	C	0.239	0.0088
More than 20%	(11 countries)	C/NC	0.149	0.0092
More than 15%	(12 countries)	NC	(0.113)	(0.0117)
More than 10%	(14 countries)	NC	(0.165)	(0.0106)

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 (2006). 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 14: Cointegration esimates by the DOLS, FMOLS and BKN methods for different samples

# countries	min weight	PDOLS	FMOLS	BKN
5	50%	0.532* (0.0032)	0.524* (0.0032)	0.284* (0.013)
6	40%	0.391* (0.0030)	0.384* (0.0029)	0.242* (0.0101)
9	30%	0.345* (0.0024)	0.336* (0.0024)	0.239* (0.0088)
11	20%	0.326* (0.0022)	0.317* (0.0021)	0.149* (0.0092)

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

Table 15: 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%.

Table 16: Break robustness check - panel unit root tests for the real exchange rates

Test	test value	block rule	crit val (5%)	block size	crit val (10%)	block size
Removal of monthly variations larger than 50%						
Test	test value	block rule	crit val (5%)	block size	crit val (10%)	block size
LLC	-6.685	SC	-8.579	68	-7.838	66
		MV	-8.960	28	-8.112	26
IPS	-2.089	SC	-2.552	47	-2.354	31
		MV	-2.620	20	-2.375	26
INVN	-0.249	SC	-2.741	115	-2.552	142
		MV	- 2.378	32	-2.122	32
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	-6.653	SC	-8.149	68	-7.516	66
		MV	-8.718	29	-8.039	26
IPS	-2.113	SC	-2.531	47	-2.344	31
		MV	-2.626	19	-2.337	28
INVN	1.273	SC	-1.450	115	-1.338	142
		MV	- 2.382	31	-2.414	26

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.

Table 17: Break robustness check - Fachin (2006)'s panel cointegration tests

	Test value	p-values basic bootstrap	FDB1	FDB2
Removal of monthly variations larger than 50%				
With common time dummies				
Mean CADF	-3.04	1.00	0.00	-3.00
Median CADF	-2.83	15.00	12.00	11.00
Without common time dummies				
Mean CADF	-3.07	3.00	4.00	5.00
Median CADF	-2.91	8.00	5.00	6.00
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	-3.08	2.00	0.00	2.00
Median CADF	-2.99	5.00	4.00	3.00
Without common time dummies				
Mean CADF	-3.07	1.00	0.00	0.00
Median CADF	-2.91	7.00	5.00	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 (2006)). P-values are in percent. See Davidson and McKinnon (2000) p.7 for comments related to the potential negativity of FDB2.

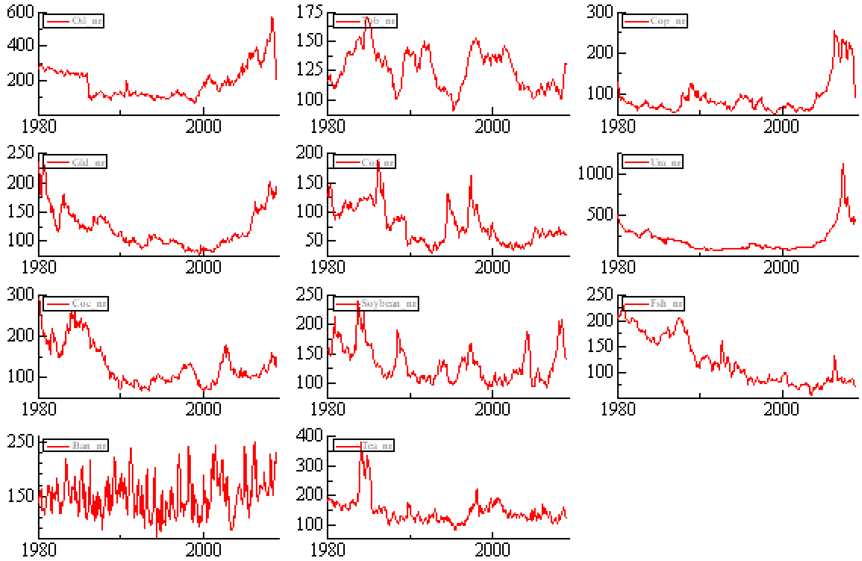


Figure 1: Selected commodity prices series, normalized (1995=100) and deflated (by the MUV) - by row, from left to right: oil, tobacco, copper, gold, coffee, uranium, cocoa, soybean, fish, banana and tea.

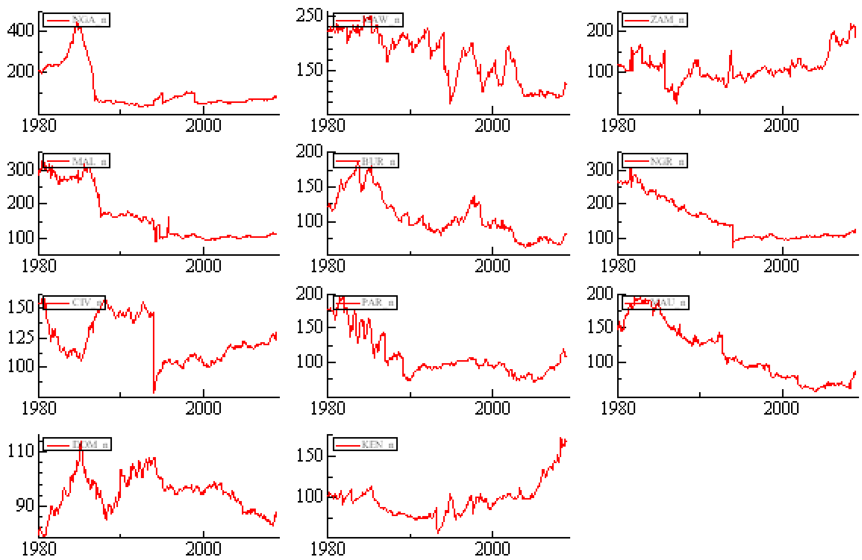


Figure 2: Normalized (1995=100) Real Effective Exchange Rates - by row, from left to right: Nigeria, Malawi, Zambia, Mali, Burundi, Niger, Côte d'Ivoire, Paraguay, Mauritania, Dominica and Kenya

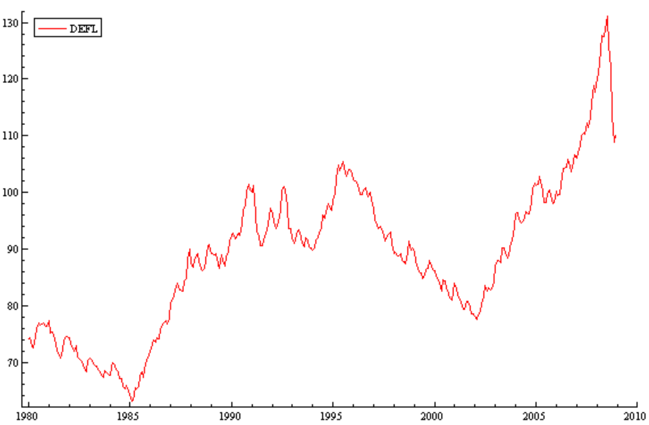
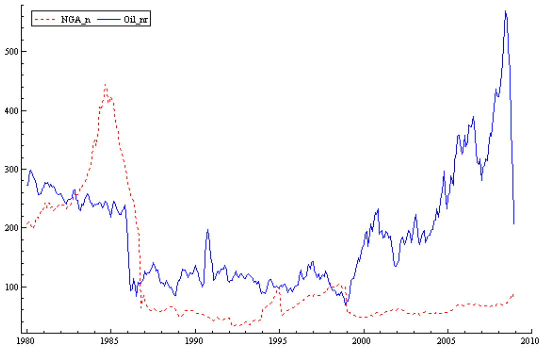
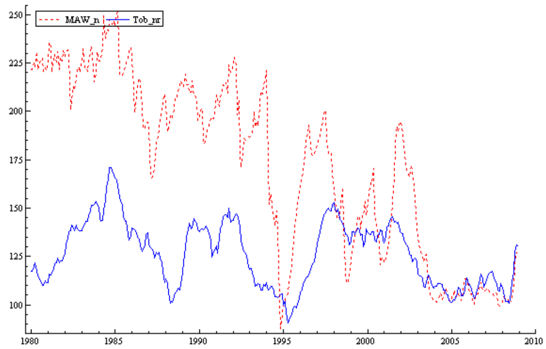


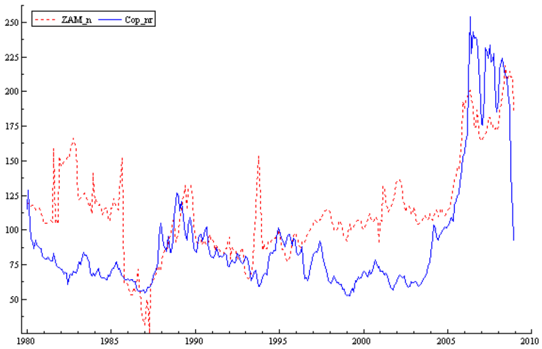
Figure 3: MUV: Unit Value Index (in US dollars) of manufactures (commodity deflator) - normalized version (1995=100)



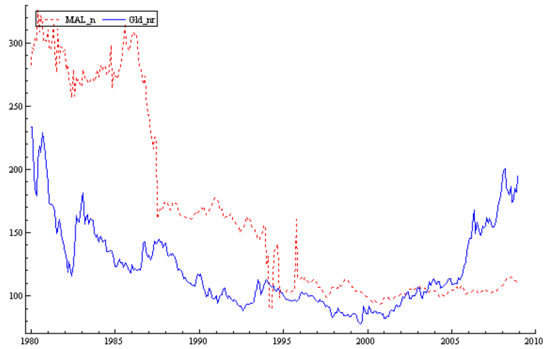
(a) Nigeria and oil



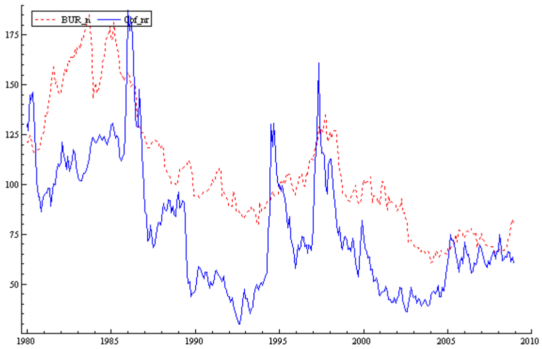
(b) Malawi and tobacco



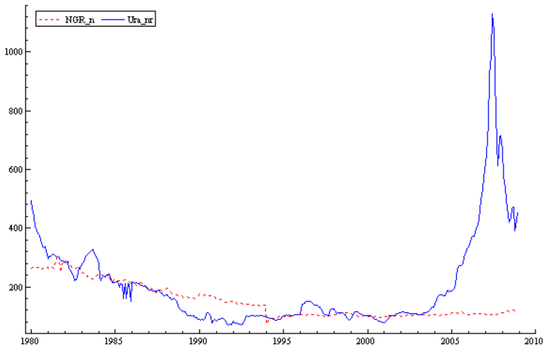
(c) Zambia and copper



(d) Mali and gold

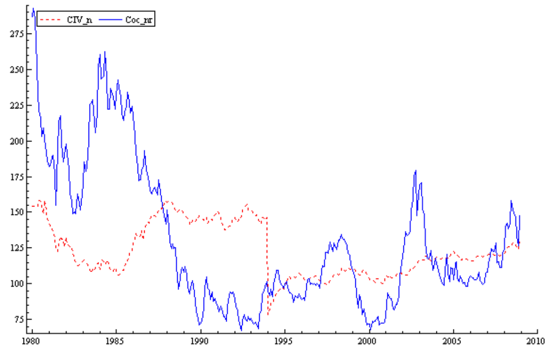


(e) Burundi and coffee

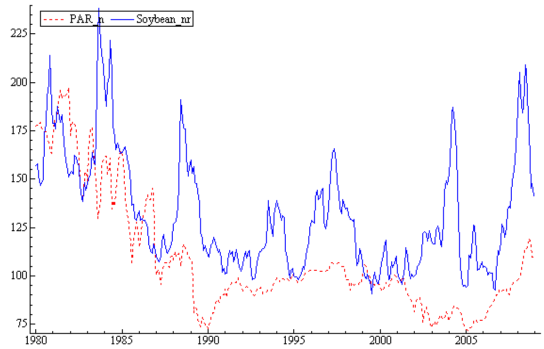


(f) Niger and uranium

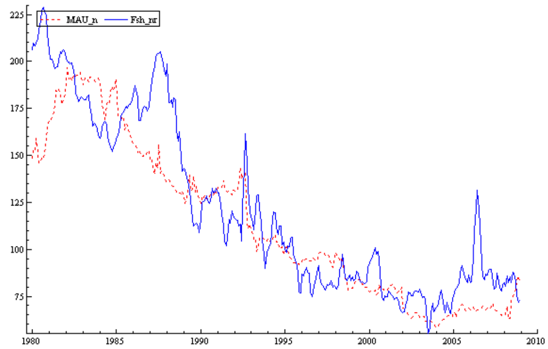
Figure 4: Real exchange rates (dashed lines) and commodity prices (solid lines). Both series are normalized (1995=100) and commodity price are deflated by the manufactures unit value index (MUV)



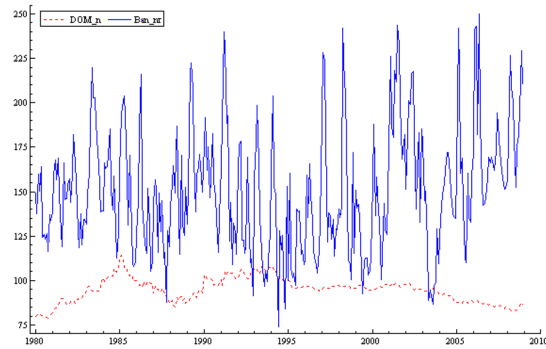
(a) Côte d'Ivoire and cocoa



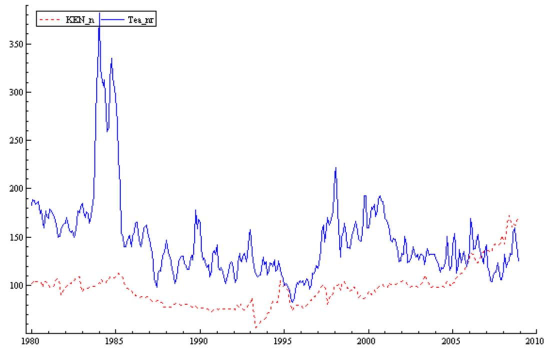
(b) Paraguay and soya



(c) Mauritania and fish



(d) Dominica and banana



(e) Kenya and tea

Figure 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)

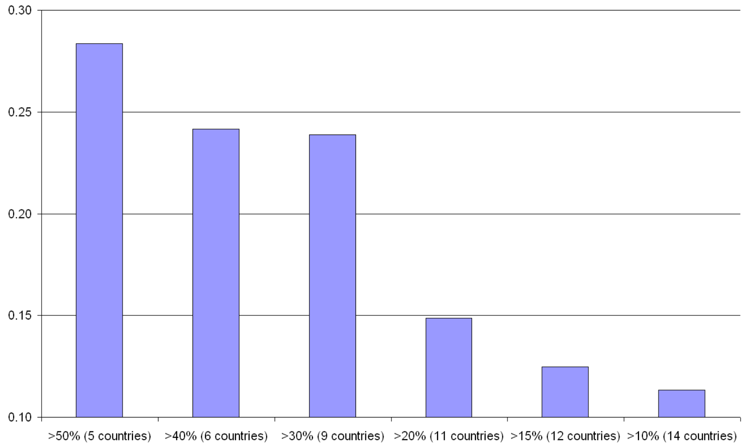


Figure 6: BKN cointegration coefficient for different samples (depending on the weight of the main commodity)

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