UNIVERSITAT DE BARCELONA

Three essays on commodity prices

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PhD in Economics

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"O frati", dissi, "che per cento milia
perigli siete giunti al l'occidente,
a questa tanto picciola vigilia

d'i nostri sensi ch'è del rimanente
non vogliate negar l'esperïenza,
di retro al sol, del mondo sanza gente.

Considerate la vostra semenza:
fatti non foste per viver come bruti,
ma per seguir virtute e canoscenza"

Dante's Inferno, Canto XXVI, versi 112-120

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Chapter 1

Introduction

1.1 The ever important relationship between commodity prices and exchange rates

Commodity prices and exchange rates tend to be connected in time and space through literature related, empirically proven linkages. From the Heckscher commodity theory to the purchasing power parity, commodities have been studied both as means of external causation for the exchange rate in small exporting countries, as well as predictors for the exchange rate movements in both the short and the long run in a more generalized framework. Past literature agrees in considering arbitrage opportunities substantial in driving fluctuations and speed of the adjustment of the conditional mean of the exchange rate through time, conditioned by the commodity prices which correlate with them in the short run; long run studies focused on estimating elasticities of the so called "behavioral definition" of the exchange rate, finding in the commodity terms of trade a valid covariate to describe it.

As Obstfeld and Taylor (1997b) pointed out, the most robust evidence on (linear) purchasing power parity is perhaps found only when long time series, spanning more than a century, are employed, as in Froot, Kim, and Rogoff (1995) in the case of the law of one price or Harvey, Kellard, Madsen, and Wohar (2010b) in the case of the Prebisch-Singer hypothesis. Traditional studies on the exchange rate and its long run structure would generally disregard any covariate correlated to it, as the existence of a non stationary process in long exchange rate series would naturally imply failure of the purchasing power parity by definition. Only in the late nineties, more than ninety years away from the seminal article of Heckscher (1916), researchers started looking more into alternative definitions guiding the long run equilibrium and short term adjustment (we would cite Obstfeld and Taylor (1997b) again, but also Imbs, Mumtaz, and Ravn (2003) and Taylor, Peel, and Sarno (2001a) for further references). As Heckscher (1916) states describing an empirical phenomenon he defined the "commodity points theory", separated commodity markets' behavior and price differentials are dictated by trade barriers. While past studies captured the effect of these barriers to real commodity trade through autoregressive studies, our attention has nowadays not only shifted in finding a behavioral,[1] thus conditional, definition for the exchange rate, but also in determining which exogenous sources (if there are any) could dominate the speed of convergence.

This thesis is aimed at testing for three pivotal aspects in the studies related to commodity prices and exchange rates: predictability, short term fluctuations and causation, and significance of the structural (long run) coefficients. Along its three chapters, this thesis will thus discuss the following points: whether or not long run estimates have changed compared to previous literature thanks to more substantial market integration in an updated sample; what would be the most correct way to shape short run fluctuations under the hypothesis of regime switching behavior, thus "resuscitating" the Heckscher commodity points theory; test for time-wise predictability of the exchange rate with all the time dimension consistent non-linear models currently at our disposal; finally, test for the specific functional form of the adjustment checking whether or not the response of the economic actors in a modern sample has become less gradual and more instantaneous (whether the slope of the transition has become more steep) thanks to modern trade liberalization and lower transactional barriers.

1.2 Research questions

This thesis analyzes the short and long run adjustment properties of a series of exchange rates in a wide set of commodity-exporting countries under the light of two

[1] See Macdonald (2000) and Macdonald (1998) for a complete overview on the behavioral definition of the exchange rate.

concurrent theories, which embody the last thirty years of applied analysis on the convergence of price differentials and exchange rates to a univariate equilibrium in observed components model analyses: the classical (linear) purchasing power parity theory; the commodity points theory, that is, the Heckscher commodity points theory seen as the studies related to the transaction cost-based interpretation of non-linearities of the real exchange adjustment path to its equilibrium.[2] As we test for the above theories, we also tested for the puzzle of exchange rate slow speed of mean reversion and consequent failure to produce reliable forecasts.[3]

Our first argument is that, contrary to the spirit of past empirical research on short run adjustment of price differentials in industrialized countries, non-linear models in emerging commodity exporting countries, as implied by the Heckscher commodity points theory, might outperform in terms of in and out-of-sample forecast the standard linear empirical models which relate to the concept of PPP convergence. Such evidence would offer a specific, country-case limited solution, to the Penn effect in commodity exporting countries. We also argue that the above hypothesis might help policy in forecasting of price differentials in order to set more efficient commercial maneuvers. To tackle such arguments, our thesis presents, in chapter two, a fundamental in-sample and out-of-sample forecasting exercise.

Our second argument has to do with the following research question: how can the long run relationship and short run deviations from a behavioral equilibrium be better captured and generalized in a panel framework? In chapter three we show how modern techniques embodying both cross sectional heterogeneity and spatial spillovers can help us depict a relationship which is arguably much less intense than what existing literature has implied in the past.[4]

Finally, our conclusive argument asks whether or not an exogenous source of variation, such as a commodity price index, might be the right tool to add to the precision of the estimates of the commodity points theory, acting as a natural threshold and

[2] Some relevant theoretical studies suggest that a smooth adjustment may be more apt at capturing transaction costs when they are proportional (as in Taylor, Peel, and Sarno (2001b)). The smooth commodity points theory can thus be defined as the aforementioned theory, only seen on the time-delaying retrospective that agents might somewhat take their time to adjust their expectations to changes in the exchange rate behavior.

[3] See Rogoff (1996) for a complete overview.

[4] More on this will be found in the informal correlation analysis in the next paragraph.

theoretically valid alternative to the absolute or relative misalignment of the exchange rate. We also argue that the shape of the transition itself has changed at the aggregated level, becoming more steep in the period we have considered when compared to previous literature estimates. This last research questions are discussed in the fourth chapter.

1.2.1 Some simple correlations

In a nutshell, our work tries to reconstruct the evolution of the relationship between commodity prices and exchange rates across time in a progressively more unified series of exercises, although without neglecting the time and space invariant heterogeneity across time and countries. It is thus worth wondering whether the relationship has, across time and with changes in preferences, changed substantially during the last four decades.

Time-wise correlation between commodities and real exchange rates has undergone periods of weakly negative persistence, followed by positive rebounds across the globe (Figure 1-1). Although being positive on average, the general agreement that a higher commodity price level contributes to the appreciation of the exchange rate through a simple increase in the tradable goods basket can not be generalized to every commodity exporter without due caveats, which across the literature were generally represented by discretional threshold related to the weight of commodity exports over the total basket of exported goods. As within country correlation shows in the graph below, heterogeneity spikes so intensely across countries[5] that no generalization appears to be possible without taking such differences under account (Figure 1-2). On top of that, and considering the average value of the within country correlation coefficient, we also argue that perhaps literature was not just overestimating the elasticity of the exchange rate with respect to commodity prices, but perhaps also failing in taking into account how short run variation patterns could be explained with transitional models, based

[5] The seventy-two countries we included, conditional on data availability, were: Australia, Austria, Belgium, Bulgaria, Bahrain, Belize, Bolivia, Brazil, Central African Republic, Canada, Switzerland, Chile, China, Cote d'Ivoire, Cameroon, Colombia, Costa Rica, Cyprus, Czech Republic, Germany, Denmark, Dominican Republic, Algeria, Spain, Finland, France, Gabon, United Kingdom, Georgia, Ghana, Gambia, Greece, Guyana, Hungary, Ireland, Iran, Iceland, Israel, Italy, Japan, Latvia, Morocco, Moldova, Mexico, Macedonia, Malta, Malaysia, Nigeria, Nicaragua, Netherlands, Norway, New Zealand, Pakistan, Philippines, Poland, Portugal, Paraguay, Romania, Russia, Saudi Arabia, Singapore, Slovak Republic, Sweden, Tunisia, Uganda, Ukraine, Uruguay, United States, Venezuela, South Africa.

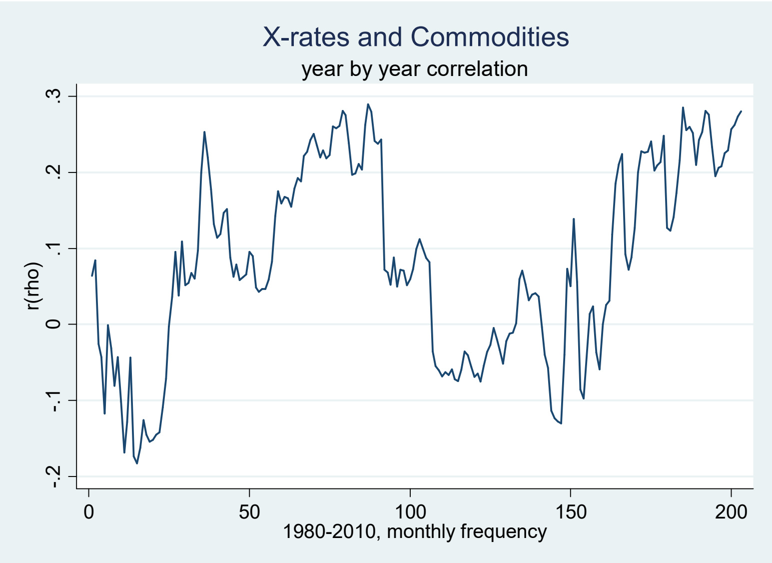

Figure 1-1: Time varying correlations of exchange rates and commodity prices

Table 1.1: Average correlation coefficients

	within time	*within countries*
$\overline{\rho}$	**0.083***	**0.186***
st.error	0,009	0.060
C.I.+	0.065	0.066
C.I.−	0.100	0.306
Points	203	70

Cross-time and cross-country correlations between the exchange rate and our commodity price index.

on speed adjustment theories which, being found consistently valid in the late nineties, could still be useful for understanding the complexity of the relationship in a more cross sectional aggregated and long span panel.

Our thesis also marks a significant contribution into the study of the exchange rate-commodities relationship by suggesting alternative means for forecasting and finally showing if Granger-causation might still be present in the relationship once old discretional thresholds and arbitrage theory converge in a single model.

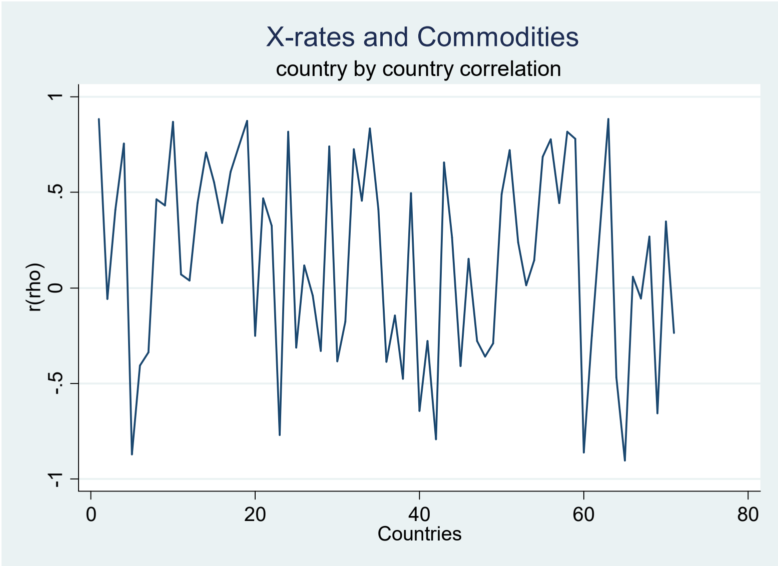

Figure 1-2: Cross sectional correlations of exchange rates and commodity prices

1.3 Contributions

Our thesis offers a series of additional contributions to the field of international financial macroeconomics, and aims, in empirical terms, at determining the modern validity of the linear and nonlinear renditions of the real exchange – commodity prices nexus. Aside from the fact of representing an update on the seminal research conducted on the two separate fields of short run adjustment conditions in country studies and general trends in aggregated ones, an additional contribution of our work stands in the field of policy making, where we argue on the possibility of a country by country anchoring of the exchange rate to the variation of the price of the most exported commodities. In the following paragraph, we shall offer an exhaustive, chapter by chapter synthesis of our main contributions.

1.3.1 Chapter two

Chapter two makes a contribution on the everlasting dispute between that branch of the literature which agrees on the existence of a linear relationship between exchange rates and commodity prices and those scholars who have focused on theoretically related regime changes in the speed of mean reversion of the exchange rate to its equilibrium with commodity prices in commodity exporting countries. As we initially focus on theoretically justifiable regime switching models, the analysis furtherly switches to analyzing the forecasting capabilities of all of them and goes beyond the call of duty by introducing and making out of sample comparisons of such models with a totally atheoretical one, namely a one layer neural network structure. To our knowledge, this exhaustive exercise represents an unprecedented attempt at comparing theoretical dependent and non-theoretical dependent modelling in this field of research, with a specific focus on out-of-sample forecasting once time-wise nonlinearity, instabilities and regime switches are taken into account in a country by country framework.

1.3.2 Chapter three

Chapter three contributes to commodity prices and exchange rate equilibria studies by adding a new degree of heterogeneity to the estimates: as previous analyses focused on framing time invariant heterogeneity through fixed effects, much of the residual variation at the cross sectional level, due to external common factors or unprecedented shocks, has been up until now almost entirely disregarded in our framework of study. As commodity markets becomes less fragmented (following a trend common to many other markets) our analysis adds to the previous literature panel estimations and looks for a common structure in the relationship between exchange rates and commodity prices by considering threshold export grouping and market maturity grouping in an exhaustive set of countries, taking into account possible cross section differences and common shocks jointly. As an additional and perhaps pivotal contribution to the exchange rate convergence theory in emerging markets, we conclude the chapter (although with due technical caveats related to the methodology) looking for Granger causation, thus offering once and for all an unprecedented and overall answer to the causation nexus between exchange rates and commodity prices which is meant to represent a conclusive addition to the seminal findings of Cashin, Céspedes, and Sahay (2004). An additional contribution of the study is moreover granted by the construction of a brand new commodity price index based on the price levels of a set of more than forty commodities in over seventy countries with fixed weights: such an updated price index is intended to work in stark contrast with the single commodity prices employed in chapter two and serves as an additional tool to investigate the existence of regime varying relationships between exchange rates and commodity prices in the long run, albeit with a major focus on short run variation.

1.3.3 Chapter four

Chapter four contributes to the studies on commodity prices and exchange rate equilibria by merging two points of view which have been treated disjointedly up until now and are embodied in chapter two and three: the former would consider single case studies in a nonlinear time series framework borrowing from previous regime-switching literature and thus focusing on time-wise nonlinear aspects of the relationship between commodity prices and exchange rates; the latter would focus on a panel data approach, considering possible heterogeneity across countries but without any inherent link to either the linear purchasing power theory or the commodity points theory. Employing the price index developed in the previous chapter, our work in the last chapter ultimately contributes in merging the two sides of the analysis presented in chapters two and three, pointing at a more generalized solution where countries grouped by export thresholds and market maturity share a common threshold behavior, thus allowing us to offer a more solid attempt at generalizing exchange rate equilibrium theories. Employing exogenous and endogenous threshold measures, we finally contribute to the exchange rate speed of adjustment puzzle literature by checking whether the magnitude of the misalignment from its long run equilibrium or an exogenous price threshold matter more or less equally in the determination of the exchange rate.

1.4 The Structure of the Thesis

Our thesis has been structured in five independent parts: the acknowledgements, an introduction, three main chapters organized as research papers and a conclusion. As the introduction and the conclusion are pretty self explanatory components of this thesis, the remaining chapters will be articulated as follows.

In chapter two, we borrow part of our framework from a seminal paper on common stochastic trends between real exchange rates and price indexes from Cashin, Céspedes, and Sahay (2004). In the chapter, we show an updated take at the country studies related to the linkage between commodities and the real exchange rate, aiming at showing how causal interpretation has been weakened by time and how more modern nonlinear models would be needed to get the most out-of-sample predictive efficiency out of the relationship.

In chapter three, we will make use of some fresh trade weights from the united nations conference for trade and development (UNCTAD) to test for the linearity and homogeneity of the structural relationship, building a new commodity price index and making use of an alternative definition of the real exchange rate. Taking into account eventual cross sectional spillovers and adding cross sectional consistent estimates to time invariant heterogeneous ones, we will show how results have changed with respect to past estimates.

In chapter four we will finally get deeper in the details of the short run adjustment to the equilibrium structure, adding panel threshold modelling to the overall picture, and showing how arbitrage opportunities and the commodity points threshold theory can still be helpful in depicting how the exchange rate behaves and adjusts itself to its conditional equilibrium in commodity exporting countries.

In the following subsections, we shall give a more complete overview of the thesis' central chapters.

1.4.1 Chapter two

As a first, physiological step in the development of our thesis, the second chapter presents a time series case study analysis of a group of small commodity exporting countries' price differentials relative to the US dollar, focusing on their non-linearity. Through threshold regression modelling we aimed at capturing the behavior of sixteen national consumers' price index (CPI) differentials relative to the US dollar CPI. Adding to the literature related to exchange rate forecasts, we calculated monthly rolling window and recursive forecasts and extended the analysis to not just threshold models but to a series of additional nonlinear models. Some form of nonlinearity is present overall in the vast majority of the analyzed countries. In particular, reconnecting ourselves to past theory on exchange rates and price differentials, we state that the parsimonious benchmark AR(1) model does not appear to perform any worse than any nonlinear model in the rolling sample exercise. The idea of a linear Purchasing Power Parity is undermined by the results of the recursive estimates and the outcome of the Diebold-Mariano type tests, which present weak evidence in favor the Heckscher commodity points theory and additionally point at alternative forms of non literature-related nonlinearity as best predictor models.

1.4.2 Chapter three

Chapter three takes the concepts from chapter two and reformulates them in a panel study, putting aside non-linearity as it was defined in the previous chapter in favor of two additional characteristics: cross sectional heterogeneity and global shocks, which would imply the existence of some degree of cross sectional dependence in the exchange rate convergence path across countries. To tackle and capture both previous characteristics in a linear setting, chapter three presents a set of estimators from the so called "mean group" family of estimators. To compare our results to past literature and further contribute to the literature on exchange rates, we estimate a behavioral real exchange rate model adopting a newly built commodity price index, which we devised starting from the most updated trade weights at our disposal. On top of that, to counterbalance the lack of suitable alternatives for controls, we choose to employ two different definitions for the real exchange rate. Our estimates show that current elasticities appear to be far lower than those estimated in past literature, especially in energy commodities exporting countries. Further Granger causality testing allows us to conclude that short term causation can be identified in the relationship between prices and exchange rates in both specialized commodity exporters and non-energy commodity exporters above a certain threshold.

1.4.3 Chapter four

Building on the result of chapter three, chapter four focuses again on non-linearity, but it does so trying to identify those factors which more likely contribute to short term fluctuations and the varying speed of adjustment of the exchange rate. We model a threshold error correction mechanism in a panel of advanced and emerging countries, selected through the use of past literature grouping criteria such as the IMF country classification as well as more literature driven discretional ways like their leading commodity export sector share over the total export value. Through threshold regression modelling, this chapter examines whether or not the absolute size of the deviation (the exchange rate variation itself), the relative behavioral misalignment (the error correction term) or commodity point-led arbitrage (the commodity price index) are responsible or not of regime-varying adjustment of the real exchange rate. Taking into account different measures of volatility in order to identify arbitrage opportunities and alternative regimes of convergence of the exchange rate to its equilibrium, we prove that the commodity points theory of Heckscher can be generalized at the longitudinal level, albeit under some (discretional and quantitative) identifying conditions, with the commodity price index acting as a reliable threshold variable.

Chapter 2

Performance of nonlinear models in international price differential forecasting in commodity exporting countries

2.1 Introduction

This chapter contributes to the existing literature on real exchange rate convergence by: focusing on price differentials behavior in a wide set of emerging countries which have been largely disregarded up until now due to data limitations. We analyze the performance of a benchmark linear model (an AR(1)) compared to a group of nonlinear models, searching for decisive conclusions on the possibility of forecasting the price differentials in commodity exporting countries, and make a final policy making suggestion. All in all, we examine and compare the behavior of sixteen currencies with respect to the US dollar and evaluate a set of nonlinear models through forecast error metrics to look for a dominant class of model and consequentially the best suited theory describing the national currency/US dollar differential.

The chapter is organized as follows: in Section 2.2, we present a brief literature review related to our analysis. In Section 2.3, we underline the theoretical framework.

In Section 2.3.1, we present our data and explain how it was treated. In Section 2.4, we present the nonlinear methods we employed for our in-sample estimates and our out-of-sample exercise. In Section 2.4.1, we illustrate the unit root tests we used as well as their results. In Section 2.4.2, we discuss the results of the F-tests on the feasibility of a nonlinear representation of the international price differentials. In Section 2.5, we present the estimated values of the attractor/error correction coefficients, calculate their half-lives, and finally compare the in-sample performance of our models with an information criterion and a measure of forecast error. In Section 2.5.1, we run a rolling window forecast over a twenty periods forecasting horizon, to compare the out of sample performance of the models to the benchmark AR(1) specification. Finally, Section 2.6 concludes the chapter.

2.2 Literature review

Nonlinearities in Exchange Rates

Literature based on non linear modelling of the exchange rates has been spanning the last thirty years of academic research. Among the many seminal contributions to that branch of international economics seeking an answer to the exchange rate slow return to its mean despite its very high (and basically intra-day) volatility, we recall the seminal work of Balke and Fomby (1997) on threshold cointegration, the fundamental contribution of Obstfeld and Taylor (1997a) on equilibrium and band threshold autoregression (equilibrium or band TAR) applications to the US CPI based price differential mean reversion in a group of six advanced countries, and lastly the important extension to the former study given by the smooth transition autoregressive application by Michael, Noobay, and Peel (1997). Although a lot has been said on price differential adjustment and exchange rate behavior in the neighborhood of its equilibrium,[1] not much attention has been given, save for a few major linear behavioral equilibrium applications,[2] to

[1] One very recent example being Allen, McAleer, Peiris, and Singh (2016b), who focused on out-of-sample forecasting performance in a series of non-linear univariate specifications for the exchange rate for a set of six hard currencies.

[2] We cite in particular Cashin, Cespedes, and Sahay (2004), who focused on a large group of commodity exporters, and Chen and Rogoff (2003a), who considered commodity exported but limited their attention to three big developed economies.

possible alternative non linear solutions to price differential forecasting in small commodity exporting countries. Our work bridges the gap between the aforementioned two research fields of international economics as it attempts to shed some light over possible nonlinearities in the speed of mean reversion of national to US CPI differentials in a set of countries specialized in exporting a single commodity with an export share of at least twenty percent over the total volume of exports.[3] Our findings suggest that the parsimonious AR(1) model used as a benchmark for our exercises does not appear to perform any worse than any nonlinear model in the rolling sample exercise. However, a set of Diebold-Mariano type tests we employed would more generally favor the Heckscher commodity points theory, but also points at alternative possible forms of nonlinearity. Given the nature of our results, and as a policy advice to small commodity exporters, we find no apparent reason to suggest commodity export price pegging as a generalized foreign exchange policy.

Recent advances in forecasting

The random walk (with drift) has always been considered the best available predictive model in terms of out-of-sample forecasting for nominal as well as real exchange rates (Meese and Rogoff (1983)). The seminar result of the former authors contradicts purchasing power parity, which would anyway have to rely on a very long time span rather than high frequencies and a large number of observations.[4] Modern forecasting is nowadays pretty much centered on the exact choice of the structure of covariates (i.e. net foreign assets and Taylor rules or a univariate approach), the frequency of the analysis and the best performing forecasting mechanism (i.e. rolling vs recursive).[5] Recently, researchers have underlined that exchange rates series could be used to predict commodity prices at quarterly frequencies,[6] while energy prices (oil) would be able to predict real exchange rates at even lower ones.[7] Perhaps the main common factor which

[3] Note that not a lot of literature has covered emerging/exporting countries yet (see again the seminal paper by Cashin, Cespedes, and Sahay (2004)). Furthermore, past literature would focus on convergence, not forecasting.

[4] See Harvey, Kellard, Madsen, and Wohar (2010a), who studied the secular decline in commodity prices for the past three centuries.

[5] See Rossi (2013).

[6] See Chen, Rogoff, and Rossi (2010).

[7] See Ferraro, Rogoff, and Rossi (2015).

unites the above authors and the main bulk of the literature on forecasting exchange rates and price differentials stands on the ground of linearity. Our paper contributes to this branch of the forecasting literature by comparing alternative nonlinear methodologies to an autoregressive benchmark in a set of commodity countries that would present nonlinearities in their price differential behavior.

Policy implications

The possibility of forecasting more efficiently price differentials makes of the real exchange rate an objective variable for policy makers. Confirmation of a price differential time behavior in line with the Heckscher commodity points theory could make the difference between a flexible policy stance or the idea of pegging the nominal exchange rate to the most exported commodity price or to a commodity price index in order to eliminate risks deriving from unwanted fluctuations. This is especially true for small exporting countries focusing on the export of a non-diversified set of goods or a single commodity. The adoption of a flexible exchange rate regime based on the exported commodity price would reap both the advantages of anchoring the exchange rate to a nominal anchor and at the same time allow for the degree of insulation from terms of trade shocks that a standard flexible exchange rate should guarantee. Among the proponents of such ideas, we mention the seminal papers of Frankel (2005a) and Frankel and Saiki (2002a).

2.3 Theoretical Framework

Our log-differential prices represent an exchange rate, defined as the cost of a basket of goods relative to the same basket between two countries, once such basket has been converted to a common numeraire. As far as the definition we employ goes, our numeraire is not a good (thus implying a real exchange rate) but a common currency. The differentials we analyze are obtained by deflating nominal indexes by the US consumer price index, expressed in dollars:

$$\frac{CPI_i}{CPI_{US}} \tag{2.1}$$

Introducing the purchasing power parity theory, and using for simplicity home and away price notation P_1 and P_2^*, the relationship between the two price baskets in levels and expressed in real terms will thus be:

$$P_1 = \varepsilon P_2^* \tag{2.2}$$

where ε represents the nominal bilateral exchange rate. The absolute version of the purchasing power parity states that the exchange rate should be equal to one, or have a tendency to return very quickly to such equilibrium in the long run. As such, for the this absolute version of the purchasing power parity theory, and making use again of our notation, we should conclude that:

$$P_1 = P_2^* \tag{2.3}$$

so that:

$$\frac{P_1}{\varepsilon P_2^*} = \frac{P_1}{P_2^*} \tag{2.4}$$

where the ratios above represent two equivalent definitions of the purchasing power parity. In principle, we would like, in a perfect textbook situation, to work with both constant values and level variables. However, given the scope of this chapter and considering our prices are time indexed, what we actually tested for is the relative version of the purchasing power parity:

$$\frac{\Delta p_{1,t}}{\Delta p_{2,t}^*} = z_t \longrightarrow k \tag{2.5}$$

for $t \longrightarrow \infty$. As the definition of relative PPP states, changes in national price levels are always equivalent to a constant value, or at least tend to such value (which does not necessarily equal the nominal exchange rate, as we saw in the absolute version) in the medium or long run. Expressing the ratio in log-deviation to take into account possible non linearities in the price index and assuming a constant rate of growth of price levels, we would get:

$$\ln(\frac{p_{1,t}}{p^*_{2,t}}) = \ln(p_{1,t}) - \ln(p^*_{2,t}) = z_t \tag{2.6}$$

Purchasing power parity can be modified introducing non-linearities in its convergence rate. The commodity point theory states that a region might exist, delimited by a lower and an upper bound on z_t, where convergence is non-existent and the price differential does not show any central tendency which would make it go back to its attractor value z_t. Such area would be due to non-perfect arbitrage conditions, which Heckscher justified with the existence of either transportation costs or uncertainty. We would as such have:

$$z = \begin{cases} \frac{\Delta p^L_{1,t}}{\Delta p^{L*}_{2,t}} \rightarrow z_1 \quad for\ t \longrightarrow \infty \\ \frac{\Delta p^M_{1,t}}{\Delta p^{M*}_{2,t}} = k, with\ k \in (z_1; z_2) \\ \frac{\Delta p^H_{1,t}}{\Delta p^{H*}_{2,t}} \rightarrow z_2\ for\ t \longrightarrow \infty \end{cases} \tag{2.7}$$

with $z_1 \neq z_2$, and where L, M and H identify the two possible states of convergence L and H, and one state of perhaps slower or non existent convergence, M. When the middle state is not present, and the threshold is reduced to $z_1 = z_2 = z$, the commodity point theory collapses to a non-linear, two regimes purchasing power parity model were price convergence might differ whether or not we are above or below the steady state value of z. Furthermore, in order to account for uncertainty and a non perfect information, the above model can be empirically adjusted with a function that smooths transition between states. Note that in this theoretical section we have not mentioned that some log run trends might be present in the price differentials. We will take such eventuality into account by detrending and demeaning z accordingly in the following sections.

2.3.1 Data

We considered an initial group of twenty-five commodity exporters. In search of a sufficiently long sample period in monthly frequencies,[8] our price differentials series were

[8] Quarterly series were available. However, as past literature has highlighted, changes in economic decisions might happen at intervals much lower than months. Perhaps, the most frequently used frequency in past works related to the topic of exchange rate behaviour has been the monthly one.

sourced from the International Monetary Fund (IMF) International Financial Statistics (IFS) database. The nominal consumer price indexes (CPI) we employ were sourced following Bodart, Candelon, and Carpantier (2015a) country selection,[9] deflated by the US CPI index, and finally demeaned and detrended according to the data generating process that was assumed from the results of a battery of unit root tests. All the variables are expressed in logarithms, and as such adjusted for eventual nonstationarity in variance, interpretable in terms of elasticities for small deviations and suitable for the simple forecasting exercise we will carry on in the chapter. Thus, as we work in log-differences, the price differentials we employ are based on the following measure:

$$z_{i,t} = \ln(CPI_{i,t}) - \ln(CPI_{US,t}) \tag{2.8}$$

where the index i represents each of the sixteen countries that were available for the analysis. The countries we consider were selected conditional on the availability of data from the source we mentioned above and a minimum export weight of their leading exported commodity close to twenty percent. A list of the selected countries is reported in Tables 2.1 and 2.2.

As we focus on the short run adjustment properties of national price differentials, being the long run trend analysis not of our main concern given the scope of this initial chapter, we will define a set of price differentials w_t as the detrended component of the price difference z_t (Nominal domestic price level over US price level) given by:[10]

$$z_t = \alpha + \beta t + w_t \tag{2.9}$$

The residuals from the above definition represent the error correction term we have used throughout the analysis. The formulation above has thus been used when no deterministic breaks could be find in the series, while the Zivot and Andrews (1992)

[9] Our selection basically entailed checking for all the available CPI index series from the IMF IFS which were employed by Bodart, Candelon, and Carpantier (2015a), conditional both on the availability of data, and on a suitable non-dollarized national currency (Ecuador was excluded from the analysis for such reason, while a CPI index series for Dominica was not available at the IMF IFS at the time of this draft).

[10] For the remainder of the chapter we will drop the subscript i for simplicity, unless reintroducing it when necessary.

Table 2.1: Country and leading commodity couples I

$COMM_t$	w_t	Weight
Cotton	Benin	61.00
	Mali	33.48
	Pakistan	20.52
Tobacco	Malawi	60.50
	Zimbabwe	19.53
Copper	Zambia	59.99
	Chile	30.79
Gold	Mali	54.05
	Burundi	35.45
	Ghana	28.56
Coffee	Burundi	50.98
	Ethiopia	46.43
	Uganda	36.87

Country and commodity couples according to Bodart, Candelon and Carpantier (2015). Highlighted in bold, the final sixteen countries group that made it to the final analysis.

Table 2.2: Country and leading commodity couples II

$COMM_t$	w_t	Weight
Uranium	Niger	41.73
	Benin	29.90
Cocoa	Ivory Coast	34.10
	Ghana	33.16
Aluminium	Mozambique	33.44
Soya	Paraguay	32.72
Fish	Mauritania	30.96
	Mozambique	19.87
Bananas	Dominica	29.20
	Ecuador	17.83
Tea	Kenya	21.20
Crustaceans	Mozambique	18.96

Country and commodity couples according to Bodart, Candelon and Carpantier (2015). Highlighted in bold, the final sixteen countries group that made it to the final analysis.

test equation has been employed when we found evidence of structural breaks. From the previous definition, the attractor/error correction term can be estimated as an OLS (ordinary least square) residual, while a constant had to be considered as data availability forced us to work with CPI, and not absolute prices. In this sense, and assuming a traditional auto-regressive process of order one, the standard work-horse formulation for convergence was:

$$\Delta\widehat{w}_t = \lambda\widehat{w}_{t-1} + e_t \tag{2.10}$$

where the error term is expected to be normally distributed with mean 0 and constant variance σ^2, with the parameter λ bounded between 0 and -2 and representing the convergence speed. As demeaning and detrending were already taken into account in the first stage OLS, the above model does not contain deterministics. Moreover, as the standard empirical formulation for the relative PPP does not contain any additional endogenous first differences, we decided to expand the models with additional lagged regressors based on a standard efficiency/parsimony trade-off analysis carried out using maximum truncation lag rules together with classic information criteria (specifically, the Bayesian information criterion). The coefficient of the error correction term w_{t-1}[11] bears important information on the efficiency of arbitrage between markets and we expect it to depend on the nature of the good(s) under consideration, the existence and type of transaction/transportation costs required to carry out the actual transaction, and other economic aspects of the activity such as the geographical distance between locations and the actual face value of the exchange rate.

Before getting to the estimates, we employ a series of unit root tests to check for three kinds of alternative hypotheses commonly accepted by literature: we employ the ADF (augmented Dickey Fuller test) to check for the benchmark linear stationary alternative; we employ the KPSS (the Kwiatkowsky, Phillips, Schmidt, and Shin (1992) stationarity test) to invert the hypotheses and consider the null of trend-stationarity, and finally we employ the ZA test (the Zivot and Andrews (1992)) to evaluate the null of a linear unit root against the alternative (and rather realistic) hypothesis of a stationary alternative

[11] Or attractor, as it has been frequently called in univariate literature: see again Obstfeld and Taylor (1997a).

with one deterministic endogenous break, possibly in both the mean and the trend.[12] This latter test, in particular, presents the following test equation:

$$\Delta z_t = \mu + \theta DU_t + \beta t + \gamma\theta DT_t + \alpha z_{t-1} + \sum_{j=1}^{k} c_j \Delta z_{t-j} + w_t \tag{2.11}$$

$$\widehat{w}_t = z_t - \left[\widehat{\mu} + \widehat{\theta} DU_t + \widehat{\beta} t + \widehat{\gamma}\widehat{\theta} DT_t + \widehat{\alpha} z_{t-1} + \sum_{j=1}^{k} \widehat{c}_j \Delta z_{t-j} \right] \tag{2.12}$$

where DU_t is a dummy variable capturing a shift in the intercept at time T_b and DT_t is a trend break variable for a break occurring at time T_b. Furthermore, $DU_t = 1$ if $t > T_b$, and zero otherwise and $DT_t = t - T_b$ if $t > T_b$ and zero otherwise. The null hypothesis is rejected if the coefficient α is statistically significant. The test is a sequential procedure, with a ten percent trimming at both ends of the series to ensure enough observations for each possible break model. Similarly to a grid-search, it makes use of a different dummy for every possible break-date. The convenient aspect of the test is that the algorithmic procedure selects the value of the break where the calculated ADF statistic (the value of $\widehat{\alpha}$) is at its minimum, thus representing the estimate that would be the most likely one to reject the unit root null.

2.4 Econometric models

Let us consider an implicit forward looking autoregressive specification of the kind:[13]

$$w_{t+s} = f(w_t, w_{t-d}, ..., w_{t-(m-1)d}; \theta) + e_{t+s} \tag{2.13}$$

where the error e_{t+s} in every given period $t, t+1, ..., t+s$ is assumed to be white noise and uncorrelated with x_{t+s} and f represents a function which maps R^m to the real numbers realm R, while θ represents the threshold value of the non-linear specification. The equation above represents variable w_t as a function of an embedded dimension (or

[12] Notice that we might have also employed an additional test, the Lumsdaine and Papell (1997), which extends the Zivot and Andrews (1992) allowing for two breaks instead of one. However, to take into account the issue of data mining and given the nature of the data, we have currently limited the analysis to just one break.

[13] For convenience we are dropping the hat notation from w.

model lags) m, a number of time delays d (delays in the transition between states) and the forecasting steps s. Considering an explicit linear representation of the above function:

$$w_{t+s} = \phi + \phi_0 w_t + \phi_1 w_{t-d}, ..., \phi_m w_{t-(m-1)d} + e_{t+s} \tag{2.14}$$

a convenient ADF reparametrization can be undertaken,

$$\Delta w_{t+s} = \phi + \lambda w_t + \zeta_1 \Delta w_{t-d} + ... + \zeta_{m-1} \Delta_m w_{t-(m-2)d} + e_{t+s} \tag{2.15}$$

where the new parameters would be: $\lambda = \phi_0 + \phi_1 + ... + \phi_m$ and $\zeta_i = -(\phi_{j+1} + \phi_{j+2} + ... + \phi_{m+j})$. The theoretical advantage of such reparametrization is that the stability conditions for all the roots of the polynomial in the original equation, the fact that they should lie all inside the unit circle, can now be easily expressed just with: $-1 < \lambda < 0$. Furthermore, on the interpretation side, it allows us to consider a mean reverting process with a short run attractor w_t, whose coefficient will represent the speed of mean reversion of the price differentials (λ). Consider now an unrestricted, three regimes first order SETAR (which we express in a non-ADF form for simplicity):

$$w_t = \begin{cases} \rho_l w_{t-1} + u_t \text{ if } w_{t-1} < \theta_l \\ \rho_m w_{t-1} + u_t \text{ if } \theta_l \leq w_{t-1} \leq \theta_h \\ \rho_h w_{t-1} + u_t \text{ if } w_{t-1} > \theta_h \end{cases} \tag{2.16}$$

the above model collapses to $w_t = \rho w_{t-1} + u_t$ whenever $\theta_l = \theta_m = \theta_h = 0$, with $\rho_l = \rho_m = \rho_h$.

In equivalent terms, the two regimes first order SETAR can be represented as:

$$w_t = \begin{cases} \rho_l w_{t-1} + u_t \text{ if } w_{t-1} \leq \theta \\ \rho_h w_{t-1} + u_t \text{ if } w_{t-1} > \theta \end{cases} \tag{2.17}$$

where the number of relevant thresholds decreases to one.[14]

[14] Notice that the above models become an X-TAR (exogenous threshold TAR models) whenever $\theta \neq w_{t-d}$. In a previous study from the author, the exogenous threshold is represented by the first order difference of the international price of petrol, and is set to be a proxy for the volatility of the energy commodity market which would affect transportation costs in international trade flows and as such the propension to operate arbitrage whenever it is convenient. The transition variable will thus not per-se be the deviation from the equilibrium path of the price differentials, but can be intepreted

All of the above models can be easily extended by adding lags in each regime[15] and, most importantly, additional restrictions can be specified to allow for $\theta_l = \theta_h$ and $\rho_l = \rho_h$, imposing two external symmetric regimes.[16] A simplified, general TAR model with p lags, d delays and m regimes will thus take the form:

$$w_t = \begin{cases} \mu_1 + \rho_{1,1}w_{t-1} + ... + \rho_{1,p_1}w_{t-p_1} + u_t \text{ if } x_{t-d} \geq \theta_{m-1} \\ \mu_2 + \rho_{2,1}w_{t-1} + ... + \rho_{2,p_2}w_{t-p_2} + u_t \text{ if } \theta_{m-1} \;\; \geq x_{t-d} \geq \theta_{m-2} \\ ... \\ \mu_m + \rho_{m,1}w_{t-1} + ... + \rho_{m,m_2}wm + u_t \text{ if } \theta_m \;\; \geq x_{t-d} \end{cases} \quad (2.18)$$

where, $x_{t-d} = w_{t-d}$ in the case of the self exciting models, and equivalent to any other selected exogenous source of variation in the case of models with an exogenous threshold.[17] In the following section, we will make use of the models discussed up until now together with a standard AR(1) model to check which one would fit better the data in an in-sample estimation exercise.

An additional non-linear model, the LSTAR (a transition autoregressive model with a logistic function as a smoothing function), together with a simple AAR (additive autoregressive model) and a NNET (linear neural network model) have also been considered and used in this chapter for an out-of-sample forecasting exercise on a forecast horizon of twenty periods, to check for consistency of the in-sample findings and to evaluate the forecasting performance of the SETAR models comparing them to alternative non-linear representations which furthermore do no support any specific economic theory. The LSTAR model can be specified as:

$$\begin{aligned} w_{t+s} \;\; = \;\; & (\phi_1 + \phi_{10}w_t + \phi_{11}w_{t-d} + \phi_{1L}w_{t-(L-1)d}(1 - G(\gamma, w_{t-d})) + \\ & (\phi 2 + \phi_{20}w_t + \phi_{21}w_{t-d} + \phi_{2L}w_{t-(H-1)d}(G(\gamma, w_{t-d})) + e_{t+s} \end{aligned} \quad (2.19)$$

where the two regimes result from a transition guided by function G, which depend on the slope coefficient γ and the delayed threshold variable w_{t-d}. The reason behind the

as one of its causes.

[15] As well as intercepts and other deterministic components.

[16] Balke and Fomby (1997) called such model a two regimes "BAND TAR" model.

[17] Notice that when $x_{t-d} = \Delta w_{t-d}$ the SETAR model becomes an M-SETAR (momentum SETAR model).

choice of a smoothing function is that no transition between two states of the world can happen discretely with market segmentation and imperfect information: as such, a portion of those who could belong to either the lower or the higher regime would basically belong to none, and be part of an "inaction band" where the series would most likely behave like a unit root. Symmetric regime applications for STAR functions have been widely used in past international economics and finance applications.[18]

An additive (non-parametric) autoregressive model was instead specified as:

$$w_{t+s} = \mu + \sum_{i=1}^{m} s_i(w_{t-(i-1)d}) \tag{2.20}$$

where d is once more the maximum delay embedded in the estimation and the s_i represent smooth functions in the form of cubic regression spines with a penalty. Finally, we estimate a simple neural network model. Such model, with D hidden units, activation function g, embedded delay d and maximum lag order of m can be represented by:

$$w_{t+s} = \beta_0 + \sum_{j=1}^{D} \beta_j g(\gamma_{0j} \sum_{i=1}^{m} \gamma_{ij} w_{t-(i-1)d}) \tag{2.21}$$

In our contribution, we refer to a simple neural network model by assuming a single (hidden) layer, that is, $D = 1$, in an attempt to mantain such non-theoretical based model as simple as possible. Our model thus reduces to:

$$w_{t+s} = \beta_0 + \beta_1 g(\gamma_{01} \sum_{i=1}^{m} \gamma_{1j} w_t) \tag{2.22}$$

so that w_t represents a unique intermediate layer result used to compute final forecasts at w_{t+s}. The last three models above have been employed only in Section 2.5.1 for out-of-sample forecasting performance comparisons.

[18] See for instance the seminal work of Taylor, Peel, and Sarno (2001b) in a univariate framework, but also the multivariate error correction estimate and subsequent Exponential STAR contribution by Baum, Barkoulas, and Caglayan (2001a).

2.4.1 Unit root analysis and demeaning

To avoid the issue of data mining we decided to couple the standard ADF and KPSS tests with the ZA unit root test with one endogenous break affecting the deterministic component of the model. The advantage of this test over the seminal Perron (1989) test is represented by the fact that break dates are endogenously retrieved by a recursive conditional estimation, and the null hypothesis is built so that the final test statistic is calculated where the model would be more likely to present a rejection of the null of unit root. The ZA test presents a trend-stationary alternative, specifically a single break in both the intercept and the trend in the alternative hypothesis. We use such test to sharpen the demeaning and detrending of z_t: non-rejection of the null hypothesis would univocally imply the presence of a unit root and exclude a break, which is accounted for in the alternative hypothesis.

The unit root tests we performed are visible in Tables 2.3 and 2.4. The standard ADF test with truncation lag order set to Schwert maximum criterion failed to reject the unit root null in all cases but one, Mozambique, while consistently with that the KPSS shows a complete rejection coverage.[19]

2.4.2 Functional form tests

In this Section we report the estimated values and bootstrapped p-values of the Hansen (1999a) F-test based on the null hypothesis of non linearity against the alternative of TAR behavior, and on the null of two regimes TAR behavior against the alternative hypothesis of three regimes. It is perhaps not surprising to see, given the extended literature related to nonlinearity in major currencies/advanced countries, that the null of linearity stands completely not rejected only in three out of the sixteen countries we analyzed, namely Chile, Paraguay and Zambia, while at least some degree of nonlinearity in the SETAR sense is detected in the remaining countries. This is visible in Tables 2.5 and 2.6, where the bootstrapped p-values of the F-test report rejection at least at

[19] To make sure the one break we choose through the endogenous selection would be at least statistically meaningful, we ran a series of ex-post Bai and Perron (2003) tests in levels or first differences, depending on the outcome of the unit root tests, on a maximum of nine stochastic (slope) breaks. The BIC criteria for the alternative break specifications of the latter test would generally indicate a maximum of two possible breaks. In more than half of our sample of countries, the first or second BP choosen break dates were close to the ones indicated by the ZA test.

Table 2.3: Unit root tests I

w_t	ADF	ZA	$Break$	$ADF/ZA\ lags$	$KPSS$	$KPSS\ lags$
(I)	(II)	(III)	(IV)	(V)	(VI)	(VII)
BEN	-2.528	-12.442***	25	6	4.218***	3
MAW	-1.149	-5.623**	175	7	2.841***	4
ZMB	-1.645	-6.833***	81	7	6.640***	4
CHL	-1.826	-4.695	48	4	2.605***	2
MAL	-2.275	-7.069***	80	7	1.990***	4
BDI	-2.275	-4.112	191	7	8.863***	4
GHA	-2.980	-5.031	182	7	8.678***	4
ETH	-0.993	-4.252	244	7	7.080***	4

Columns (2) and (3): ADF and ZA t-statistics. Columns (5) and (7) lag truncation orders selected according to Schwert minimal criterion. Column (4): endogenously retrieved period of deterministic break. Column (6): KPSS test statistic.

Table 2.4: Unit root tests II

w_t	ADF	ZA	$Break$	$ADF/ZA\ lags$	$KPSS$	$KPSS\ lags$
(I)	(II)	(III)	(IV)	(V)	(VI)	(VII)
UGA	-1.515	-4.086	105	6	6.843***	3
NGR	-1.630	-7.377***	168	7	1.598***	4
CIV	-1.915	-8.052***	168	7	7.791***	4
MOZ	-4.142***	-6.374***	18	6	6.344***	3
PAR	0.336	-5.554**	113	7	8.289***	4
MAU	-2.468	-3.854	228	7	7.575***	4
DOM	-2.617	-3.766	231	7	7.984***	4
KEN	-1.285	-4.946*	146	7	8.776**	4

Columns (2) and (3): ADF and ZA t-statistics. Columns (5) and (7) lag truncation orders selected according to Schwert minimal criterion. Column (4): endogenously retrieved period of deterministic break. Column (6): KPSS test statistic.

Table 2.5: Hansen linearity test I

w_t	$F-test$	$p-value$	$F-test$	$p-value$	$F-test$	$p-value$
(I)	(II)	(III)	(IV)	(V)	(VI)	(VII)
BEN	35.328**	0.045	95.666*	0.100	53.845**	0.030
MAW	35.776*	0.060	44.398	0.254	7.966	0.690
ZMB	31.116	0.111	62.246	0.170	28.742	0.450
CHL	12.726	0.365	19.023	0.788	5.517	0.940
MAL	49.604**	0.019	91.437**	0.023	36.575**	0.120
BDI	21.068	0.151	48.377**	0.030	26.425**	0.020
GHA	62.877***	0.005	108.676***	0.007	40.015**	0.050
ETH	34.436***	0.001	55.714***	0.002	19.718	0.205

Columns (2) (4) and (6):Hansen F-test values respectively testing one regime against two, one regime against three and two regimes against three. Columns (3), (5) and (7): bootstrapped p-value (currently 1000 iterations).

the ten percent significance level in the majority of the countries for either the F-test alternative of a three regimes SETAR model or a two regimes SETAR model.

Although the F-tests results visible in Columns (II) to (V) of Tables 2.5 and 2.6 represent a valid way to infer on the functional form of the error correction mechanism, this ex-ante value is not indicative of whether or not we should choose a specific functional form over another. As a matter of fact, Tables 2.5 and 2.6 come with two additional columns, VI and VII, where the F-tests for the null of two regimes against an alternative specification of a BAND-TAR with two non-symmetric thresholds is evaluated for each country. The columns presents mixed results: none of the two models clearly prevails over the other. In order to give better inference and be more conclusive about the in-sample goodness of fit of the models, we present in Section 2.5, the estimates, the in-sample Mean Absolute Percentage Error and the Bayesian Information Criterion for all the computed SETAR models plus the AR(1) benchmark.

2.5 Estimates and in-sample fitness

Estimates of the AR(1), two regimes SETAR and unrestricted three regimes SETAR are reported in Tables 2.7 and 2.8. The AR(1), two regime and three regime SETAR

Table 2.6: Hansen linearity test II

w_t	$F-test$	$p-value$	$F-test$	$p-value$	$F-test$	$p-value$
(I)	(II)	(III)	(IV)	(V)	(VI)	(VII)
UGA	23.352**	0.035	46.728**	0.017	21.634*	0.079
NGR	39.815**	0.014	72.299**	0.018	29.760*	0.075
CIV	23.152	0.142	59.485*	0.083	34.497*	0.090
MOZ	85.061***	0.010	273.356***	0.00	143.688***	0.010
PAR	20.600	0.213	40.310	0.348	18.819	0.370
MAU	48.570***	0.000	84.195***	0.00	31.481***	0.000
DOM	26.414**	0.022	48.204**	0.025	20.542	0.170
KEN	35.380*	0.054	42.207	0.288	6.313	0.870

Columns (2) (4) and (6):Hansen F-test values respectively testing one regime against two, one regime against three and two regimes against three. Columns (3), (5) and (7): bootstrapped p-value (currently 1000 iterations).

would refer respectively to the classic PPP theory, an asymmetric variation of it[20] and the revised Heckscher commodity points theory with an inaction band. As we would expect, the magnitude of the various speed of adjustment coefficients λ greatly varies not just across countries but also across price regimes.

We would generally find no absolute confirmation of the PPP nor of the Heckscher commodity points theory. Basing our inference on both the results of the linearity tests and the in-sample fitness of our models. Failure to reject linearity would represent a valid argument in favor of the nonlinear alternative, which would in turn give us some degree of confirmation on the commodity points theory based on the validity of the estimates of the threshold models employed in the analysis. In general, all countries would seem to find the attractor for the AR(1) process significant, exception done for Uganda, Chile and Ethiopia. In most cases, the SETAR three regimes model, which we considered as the ideal setup for representing the inaction band typically found in the advanced countries literature, showed significant unrestricted coefficients in the outer regimes and an inner unit root only in Ghana, Niger, Ivory Coast and Kenya. All in all, no specification appears to be significantly more suited than another, at least in terms of suggesting a prevailing economic theory.

[20] Depending on which between the home and the away (US) currency is currently devaluated with respect to the other. In a sense, it is just a less empirically valid rendition of the Heckscher commodity points theory.

In Table 2.9, we compare the implied half-lives of the deviations of the price differentials across regimes and models. As we would expect, the autoregressive estimates would indicate some of the half-lives ranging from three to more than six years. Strikingly, in nine of the countries we considered, deviations in the AR(1) specifications do not survive the one year. This is still high considering the volatility of the currency market, but perhaps lower than what implied by Rogoff (1996). The two-regime estimates do not generally confirm faster speed of mean reversion, and in some instances, quite puzzlingly, they show slower adjustment compared to the linear alternative, in both the low and the high regime.

As we finally turn to the in-sample fitness of the models, we see that non linear modelling appears to outperform the benchmark linear specification in five out of the sixteen analyzed countries. We choose to evaluate the in-sample fitness using the Bayesian information criteria, whose results are visible in Table 2.10. Results did not differ that much when we used the Aikake information criterion and the mean absolute percentage errors, which are visible in Table 2.11.

In the upcoming Section 2.5.1 we will present our forecasting exercise and employ the MAPE once more to check how the alternative model specifications would perform outside of the sample.

2.5.1 Out-of-sample results

In our forecasting exercise, we apply a rolling window procedure over the first two-hundred and fifty-three consecutive periods (up to January 2011) for a forecast horizon of twenty months. As we have stated in the introduction of this chapter, and considering the mixed results obtained in the in-sample performance by the non linear models adopted, it seemed logical to extend the exercise adding a second subgroup of models representing additional forms of nonlinearity in our price differentials, composed of an LSTAR with two regimes, a threshold delay of two and an amount of lags depending on Schwert minimal length criteria, an AAR model and finally a LNN model with similar lag orders. As we performed the exercise on a group of countries with a non standard series length, we generally had enough observations to let the forecasting window remain fixed in relationship to the time series span, but had to modify it in the

Table 2.7: Non linear estimates I

	$AR(1)$	$SETAR$ 2 $regimes$		$SETAR$ 3 $regimes$		
w_t	λ	λ_l	λ_h	λ_l	λ_m	λ_h
(I)	(II)	(III)	(IV)	(V)	(VI)	(VII)
BEN	-0.222***	0.038	-0.161***	0.038	-0.191*	-0.093
	('0.037)	(0.067)	(0.059)	(0.067)	(0.097)	(0.079)
MAW	-0.050***	-0.062***	-0.023	-0.046**	-0.081*	-0.037
	(0.015)	(0.019)	(0.023)	(0.022)	(0.048)	(0.023)
ZMB	-0.037***	-0.022	-0.028	-0.017	-0.361***	-0.003
	(0.014)	(0.020)	(0.027)	(0.022)	(0.108)	(0.027)
CHL	-0.071	-0.040	-0.026	-0.040	0.001	-0.061
	(0.022)	(0.034)	(0.029)	(0.034)	(0.051)	(0.038)
MAL	0.131***	-0.087**	-0.059	-0.078*	-0.208***	-0.033
	(0.026)	(0.036)	(0.047)	(0.043)	(0.061)	(0.048)
BDI	-0.017**	-0.003	-0.026**	-0.006	0.007	-0.041**
	(0.007)	(0.09)	(0.010)	(0.008)	(0.024)	(0.013)
GHA	-0.026***	-0.039***	-0.014	-0.039***	0.007	-0.018*
	(0.007)	(0.09)	(0.009)	(0.009)	(0.031)	(0.009)
ETH	-0.002	-0.004	-0.023	-0.04	-0.008	0.001
	(0.003)	(0.005)	(0.005)	(0.005)	(0.012)	(0.006)

Column (1): AR(1) error correction term estimates, S.E. in parenthesis. Columns (2) and (3): Error correction coefficient values for the self-exciting TAR with two regimes and one threshold specification. Columns (5), (6) and (7): Error correction coefficient values for the self-exciting TAR with 3 regimes and two thresholds specification.

Table 2.8: Non linear estimates II

	$AR(1)$	$SETAR$ 2 $regimes$		$SETAR$ 3 $regimes$		
w_t	λ	λ_l	λ_h	λ_l	λ_m	λ_h
(I)	(II)	(III)	(IV)	(V)	(VI)	(VII)
UGA	-0.008	-0.000	-0.036**	-0.015	(0.016)	-0.036**
	(0.008)	(0.008)	(0.014)	(0.011)	(0.012)	(0.014)
NGR	-0.087***	-0.100***	-0.024	-0.105***	-0.046	-0.071**
	(0.017)	(0.024)	(0.036)	(0.024)	(0.066)	(0.034)
CIV	-0.099***	-0.133***	-0.045	-0.129***	0.048	-0.050*
	(0.021)	(0.028)	(0.031)	(0.028)	(0.285)	(0.031)
MOZ	-0.071**	-0.119***	0.033	-0.118***	-0.172	-0.040
	(0.022)	(0.025)	(0.039)	(0.021)	(0.085)	(0.037)
PAR	-0.031***	-0.022	-0.037*	-0.014	-0.134**	-0.033
	(0.012)	(0.524)	(0.021)	(0.015)	(0.055)	(0.022)
MAU	-0.074***	-0.023	-0.155***	-0.018	0.006	-0.231***
	(0.021)	(0.027)	(0.034)	(0.031)	(0.038)	(0.049)
DOM	-0.078***	-0.050**	-0.054**	-0.036	-0.035	-0.044
	(0.017)	(0.024)	(0.025)	(0.024)	(0.075)	(0.026)
KEN	-0.064***	-0.034	-0.127***	-0.060**	0.023	-0.100***
	(0.017)	(0.022)	(0.032)	(0.030)	(0.036)	(0.031)

Column (1): AR(1) error correction term estimates, S.E. in parenthesis. Columns (2) and (3): Error correction coefficient values for the self-exciting TAR with two regimes and one threshold specification. Columns (5), (6) and (7): Error correction coefficient values for the self-exciting TAR with 3 regimes and two thresholds specification.

Table 2.9: Half-lives estimates, AR(1) and best SETAR

	$AR(1)$	$SETAR\ (BEST\ BIC)$		
w_t	λ_{months}	$\lambda_{l,months}$	$\lambda_{m,months}$	$\lambda_{h,months}$
(I)	(III)	(IV)	(V)	(VI)
BEN	2.757***	0	-	3.949***
MAW	13.555***	10.830***	-	29.790
ZMB	18.284***	31.159	-	24.407
CHL	9.379	16.980	-	26.311
MAL	4.948***	7.615**	-	11.398
BDI	40.997**	230.702	-	26.311**
GHA	26.595***	17.424***	98.674	38.161*
ETH	279.826	172.940	-	29.789
UGA	88.381	Inf	-	18.905**
NGR	7.588***	6.579***	-	28.533
CIV	6.683***	4.857***	-	15.054
MOZ	9.363**	5.520***	3.672	16.980
PAR	21.832***	31.159	-	18.385*
MAU	9.085***	29.789	-	4.116***
DOM	8.572***	13.513***	-	12.486***
KEN	10.515***	20.038	-	5.103***

Half-life estimates of the attractors/error correction parameters, calculated according to $\lambda^T = (1-x)/(1-\lambda)$. The value of x represents the set decaying period, 0.5, indicating the magnitude of readjustment for which T is calculated. Columns (4), (V) and (VI): estimates for the best BIC performing SETAR model.

Table 2.10: BIC results on model fitness

w_t	$AR(1)$	$SETAR$ 2 *regimes*	$SETAR$ 3 *regimes*
(I)	(II)	(III)	(IV)
BEN	-2175.675	-2173.325	-2144.747
MAW	-2598.134	-2577.289	-2565.858
ZMB	-1653.474	-1617.524	-1598.541
CHL	-1055.621	-1039.028	-1028.198
MAL	-2735.672	-2735.852	-2712.456
BDI	-3426.280	-3403.385	-3373.850
GHA	-3215.846	-3228.632	-3239.407
ETH	-3383.781	-3393.565	-3361.018
UGA	-2633.685	-2615.080	-2604.004
NGR	-3217.701	-3214.032	-3189.330
CIV	-3614.405	-3607.533	-3563.952
MOZ	-1902.421	-1923.465	-2005.764
PAR	-3036.307	-3027.375	-2992.899
MAU	-3236.322	-3234.195	-3224.534
DOM	-4169.419	-4148.434	-4127.868
KEN	-3173.778	-3139.624	-3126.026

Columns (2) to (4): Bayesian Information criterion results for the in-sample estimates across all models.

Table 2.11: MAPE results on model fitness

w_t	$AR(1)$	$SETAR$ 2 *regimes*	$SETAR$ 3 *regimes*
(I)	(II)	(III)	(IV)
BEN	2.492%	2.695%	2.915%
MAW	2.088%	2.303%	2.130%
ZMB	6.283%	11.621%	4.626%
CHL	1.411%	1.492%	1.539%
MAL	1.726%	2.286%	2.253%
BDI	2.099%	2.498%	1.935%
GHA	2.650%	2.960%	3.047%
ETH	2.320%	2.279%	4.059%
UGA	1.969%	1.983%	2.473%
NGR	2.233%	2.695%	2.532%
CIV	1.683%	1.378%	1.551%
MOZ	3.983%	10.429%	4.715%
PAR	2.542%	2.841%	2.668%
MAU	1.820%	2.883%	1.802%
DOM	1.544%	1.914%	2.007%
KEN	2.461%	2.472%	2.657%

Columns (2) to (4): Mean average percentage error results for the in-sample estimates across all models.

case of Mozambique, whose time series presented less than one hundred observations. Error forecast measurements for the forecasting exercise are available in Table 2.12. The value for the SETAR belongs to the model which showed the lowest information criteria in the in-sample exercise. We remind the reader that the measure we are using is defined as:

$$MAPE = \frac{100}{T} \sum_{t=1}^{T} \left| \frac{w_t - \widehat{w}_t}{w_t} \right| \tag{2.23}$$

where w_t would be the actual value, $\widehat{w}_t$ the forecast, and $t \in (1; T)$ the forecasting horizon. Results of this exercise show how lowest percentage errors appear to be evenly distributed across all but the self-exciting TAR models, which perform badly across all countries. The results for the percentage error, exception done for Ethiopia, appear to be relatively high and cast a shadow over asymmetrical TAR models as a possible alternative to univariate piecewise linear models in price differential forecasting. Finally, the best relative performance can be attributed to the AR(1) model and the NNET model. In order to check this findings, and considering the drawbacks of the percentage deviation measure we employed,[21] we calculated an additional measure, namely the root mean square error (RMSE), where the divergence from the actual value would be standardized by the number of forecasting periods:

$$RMSE = \sqrt{\frac{\sum_{t=1}^{T} (\widehat{w}_t - w_t)^2}{T}} \tag{2.24}$$

Results for this additional measure are reported in Table 2.13.

Although with slight differences due to the approximation of the measures, the lowest RMSE appears to be evenly distributed across the various models, with the AR(1), the NNET and the AAR generally contending the crown of best performer. Once more, the LSTAR model appears to contribute relatively less to the out-of-sample forecast performance, while the best BIC-selected SETAR model would never show a satisfying performance. To ultimately cope with the somewhat more obscure interpretation of the RMSE, we also calcuated the mean absolute error (MAE), which computes the average

[21] Among all, possible 0 divisions when real values are to close to such limit, which happens to be the case in some instances of our differential measure.

Table 2.12: Out-of-sample forecasting performance MAPE

w_t	$AR(1)$	$SETAR$	$LSTAR$	$NNET$	AAR
(I)	(II)	(III)	(IV)	(V)	(VI)
BEN	141.332%	2189.900%	164.154%	151.376%	184.235%
MAW	64.340%	108.115%	46.716%	38.425%	40.706%
ZMB	25.959%	49.676%	31.861%	35.202%	25.345%
CHL	66.248%	530.981%	63.153%	61.412%	64.499%
MAL	103.397%	1470.881%	140.889%	109.904%	116.445%
BDI	309.489%	945.915%	294.758%	285.606%	290.494%
GHA	115.388%	128.156%	131.857%	134.527%	132.761%
ETH	7.027%	92.449%	7.308%	6.956%	7.524%
UGA	25.471%	109.531%	24.781%	23.473%	24.347%
NGR	225.867%	3591.119%	125.791%	150.204%	137.656%
CIV	92.460%	1077.125%	95.737%	93.782%	98.953%
MOZ	29.907%	361.428%	34.514%	19.998%	31.154%
PAR	41.697%	59.775%	62.103%	62.556%	46.763%
MAU	38.535%	129.570%	45.906%	40.335%	41.548%
DOM	339.129%	4920.560%	464.246%	436.971%	384.256%
KEN	95.013%	736.287%	87.591%	105.982%	111.342%

Columns (2) to (6): Mean average percentage error for all the models employed in the analysis.

Table 2.13: Out-of-sample forecasting performance RMSE, rolling

w_t	$AR(1)$	$SETAR$	$LSTAR$	$NNET$	AAR
(I)	(II)	(III)	(IV)	(V)	(VI)
BEN	0.0172	0.1462	0.0231	0.0169	0.0206
MAW	0.0400	0.0869	0.0357	0.0362	0.0356
ZMB	0.0604	0.1556	0.0673	0.0694	0.0597
CHL	0.0039	0.0303	0.0036	0.0036	0.0035
MAL	0.0096	0.0912	0.0111	0.0101	0.0101
BDI	0.0148	0.0737	0.0147	0.0606	0.0147
GHA	0.0171	0.0416	0.0199	0.0200	0.0193
ETH	0.0445	0.3766	0.0453	0.0448	0.0462
UGA	0.0208	0.1069	0.0212	0.0209	0.0209
NGR	0.0112	0.0971	0.0109	0.0109	0.0111
CIV	0.0101	0.0601	0.0107	0.0104	0.0108
MOZ	0.0160	0.1038	0.0192	0.0146	0.0167
PAR	0.0307	0.0537	0.0333	0.0344	0.0308
MAU	0.0177	0.0521	0.0192	0.0197	0.0192
DOM	0.0064	0.0712	0.0067	0.0065	0.0068
KEN	0.0097	0.0646	0.0097	0.0097	0.0098

Column (2) to (6): Root mean square error for all the models employed in the analysis.

Table 2.14: Out-of-sample forecasting performance MAE

w_t	$AR(1)$	$SETAR$	$LSTAR$	$NNET$	AAR
(I)	(II)	(III)	(IV)	(V)	(VI)
BEN	0.011	0.146	0.013	0.010	0.013
MAW	0.032	0.074	0.025	0.023	0.024
ZMB	0.025	0.141	0.022	0.022	0.021
CHL	0.003	0.029	0.003	0.003	0.003
MAL	0.008	0.089	0.009	0.008	0.008
BDI	0.010	0.062	0.011	0.011	0.011
GHA	0.013	0.036	0.012	0.012	0.012
ETH	0.014	0.370	0.015	0.014	0.015
UGA	0.009	0.101	0.010	0.009	0.009
NGR	0.009	0.100	0.008	0.008	0.008
CIV	0.007	0.057	0.007	0.007	0.007
MOZ	0.011	0.087	0.013	0.010	0.010
PAR	0.011	0.044	0.011	0.012	0.011
MAU	0.007	0.039	0.008	0.008	0.008
DOM	0.005	0.070	0.005	0.005	0.005
KEN	0.006	0.053	0.006	0.006	0.006

Columns (2) to (6): Mean absolute Error for all the models employed in the analysis.

absolute difference between $\widehat{w}_t$ and w_t and allows proportional contribution of each error to the absolute measure. The MAE was calculated as:

$$MAE = \sum_{t=1}^{T} \left| \frac{\widehat{w}_t - w_t}{T} \right| \tag{2.25}$$

results are available in Table 2.14, and the conclusions we could draw from these results did not differ from those previously seen with other forecast deviation measures. One additional insight from this more direct measure of forecast error is that, perhaps in line with what we would have normally expected from the threshold models, the NNET and AAR models appear to statistically dominate the residual models even more, with eleven and twelve of the best relative results out of a total of sixteen countries respectively.

To conclude our forecasting exercise, we present the recursive one-step-ahead forecast estimates of our models (Table 2.15) and the Diebold-Mariano tests (Table 2.16) evaluating any systematic difference between the linear autoregressive specification across the two different forecasting methods, aimed at helping us in the ex-post choice of the optimal forecast.

The rolling forecasts would suggest the SETAR models as an alternative as valid as (if not slightly better than) the linear specification for out-of-sample forecasting. This would represent weak evidence pushing our conclusion in favor of the Heckscher commodity points theory rather than the classic purchasing power parity theory. However, we ultimately choose to specify a set of two Diebold-Mariano type tests, in order to compare the autoregressive specification in both the rolling and the recursive estimates. In such instance we have conclusive evidence over the better performance of the AR(1) in the rolling forecast exercise compared to the recursive estimates. When considering forecasting purposes, the simple and more parsimonious AR(1) model does not appear to perform any worse than any nonlinear model in the out of sample rolling-window forecasting exercise. In theoretical terms, the AR(1) model still appears to outperform any possible theory-related nonlinear variant of the relative price adjustment in rolling window forecasting, but its interpretation as a long run equilibrium definition is overshadowed by the results of the recursive estimates, which present a systematically lower RMSE and should theoretically guarantee relatively lower error forecasts given that the procedure fixes the beginning of the forecast window at the very beginning of the sample.

2.6 Conclusions

In this chapter, we presented an analysis of sixteen commodity countries' exchange rate movements in relationship to the US dollar. In-sample fitness of sixteen US CPI-relative national price differentials is evaluated and modelled using a set of threshold nonlinear models, while out-of sample forecast accuracy is evaluated on the basis of rolling and fixed window forecasts through calculation of mean absolute percentage errors and extended to three additional models, namely a logistic smooth transition regression, an additive non linear autoregressive model and a simple neural network model. Our preliminary results confirm presence of a form of TAR non linearity in the majority of the countries analyzed. SETAR models tend to have quite poor relative performance in the rolling window exercise, both when compared to alternative nonlinear specifications and to the benchmark linear model, but perform at least as well as the linear specification in the recursive exercise. Although this pushes our conclusions towards the acceptance

Table 2.15: Out-of-sample forecasting performance RMSE, recursive

w_t	$AR(1)$	$SETAR$	$LSTAR$	$NNET$	AAR
(I)	(II)	(III)	(IV)	(V)	(VI)
BEN	0.0393	0.1290	0.1061	0.0652	0.0571
MAW	0.2989	0.1817	0.2730	0.2826	0.2711
ZMB	0.3455	0.1085	0.3511	0.3320	0.3629
CHL	0.0145	0.0612	0.0263	0.0160	0.0154
MAL	0.0134	0.1119	0.0148	0.0201	0.0167
BDI	0.0810	0.0643	0.0663	0.0746	0.0722
GHA	0.0532	0.0374	0.0509	0.0774	0.0628
ETH	0.2095	0.2770	0.2102	0.2295	0.2045
UGA	0.1377	0.0347	0.1304	0.1312	0.1316
NGR	0.0222	0.1226	0.0284	0.0322	0.0277
CIV	0.0362	0.1246	0.0479	0.0425	0.0534
MOZ	0.0722	0.1975	0.1287	0.0409	0.0391
PAR	0.1834	0.0588	0.1936	0.2084	0.1823
MAU	0.0926	0.0221	0.1118	0.5696	0.1227
DOM	0.0632	-	0.0330	0.0305	0.0134
KEN	0.0901	0.1141	0.0878	0.0916	0.1067

Column (2) to (6): Root mean square error for all the models employed in the analysis.

Table 2.16: Diebold-Mariano tests

w_t	DM^2	$p-value$	DM	$p-value$	$AR(1)^{rolling}$	$AR(1)^{recursive}$
(I)	(II)	(III)	(IV)	(V)	(VI)	(VII)
BEN	1.8509	0.0912	2.9474	0.0133	0.0172	0.0393
MAW	8.1317	0.0000	11.0460	0.0000	0.0400	0.2989
ZMB	6.8603	0.0000	17.5900	0.0000	0.0604	0.3455
CHL	0.9647	0.3554	1.3566	0.2021	0.0039	0.0145
MAL	-0.4503	0.6612	0.0152	0.9882	0.0096	0.0134
BDI	3.8439	0.0027	5.8825	0.0001	0.0148	0.0810
GHA	1.7745	0.1036	2.1085	0.0587	0.0171	0.0532
ETH	6.0500	0.0001	43.984	0.0000	0.0445	0.2095
UGA	10.4860	0.0000	114.13	0.0000	0.0208	0.1377
NGR	1.8199	0.0961	1.8154	0.0968	0.0112	0.0222
CIV	2.2726	0.0441	3.1910	0.0086	0.0101	0.0362
MOZ	6.8556	0.0000	154.5800	0.0000	0.0160	0.0722
PAR	8.3821	0.0000	27.5550	0.0000	0.0307	0.1834
MAU	6.6221	0.0000	14.2220	0.0000	0.0177	0.0926
DOM	18.2240	0.0000	39.0550	0.0000	0.0064	0.0632
KEN	5.8573	0.0001	5.7841	0.0001	0.0097	0.0901

Column (2) and (4): Test statistic based on squared and linear absolute difference loss function between the previous rolling and recursive raw forecasts.

Table 2.17: Descriptive statistics, price differentials

w_t	$Min.$	$1st\ Q.$	$Median$	$Mean$	$3rd\ Q.$	$Max.$
(I)	(II)	(III)	(IV)	(V)	(VI)	(VII)
BEN	-0.389	-0.013	0.005	0.000	0.024	0.078
MAW	-0.840	-0.095	-0.007	0.000	0.109	0.367
ZMB	-1.982	-0.194	0.101	0.000	0.281	0.469
CHL	-0.030	-0.016	0.003	0.000	0.012	0.058
MAL	-0.281	-0.023	0.005	0.000	0.025	0.102
BDI	-0.393	-0.078	0.026	0.000	0.095	0.374
GHA	-0.682	-0.110	0.019	0.000	0.130	0.492
ETH	-0.540	-0.205	0.033	0.000	0.193	0.488
UGA	-0.170	-0.110	0.020	0.000	0.094	0.139
NGR	-0.367	-0.025	0.007	0.000	0.035	0.194
CIV	-0.295	-0.017	0.001	0.000	0.017	0.100
MOZ	-0.526	-0.048	0.024	0.000	0.062	0.119
PAR	-0.528	-0.107	0.050	0.000	0.100	0.238
MAU	-0.095	-0.024	0.010	0.000	0.025	0.111
DOM	-0.078	-0.016	0.002	0.000	0.017	0.068
KEN	-0.386	-0.042	0.011	0.000	0.048	0.231

Descriptive statistics for every series. The variables have been demeaned (see Column (5)) and detrended.

of the commodity points theory, the results obtained with the additive autoregressive model and the neural network suggest further investigation into alternative nonlinear forecasting methods which are currently unrelated to the theory which has been developed up until now. Our results finally confirm the presence, in the spirit of Obstfeld and Taylor (1997a) review of the Heckscher commodity points theory, of band convergence with an inner unit root in seven of the sixteen countries considered once the final recursive forecasting estimates are accounted for. As a final policy advice to commodity exporting countries, we find no apparent reason to suggest commodity export price pegging as a generalized policy perhaps given the presence of strong nominal frictions dampening relative price adjustment.

Chapter 3

Long run elasticities and short run adjustment of real exchange rates in commodity exporting countries

3.1 Introduction

This chapter contributes to the existing literature on real exchange rate convergence by employing "macro-panel" estimators in a panel set-up where time dimension is high compared to cross sectional dimension, allowing for slope heterogeneity to affect the study of commodity prices comovement with the exchange rate in the long run. Specifically, we analyze the impact of commodity prices on the sluggish real exchange rate convergence rate with a monthly time frequency and monitor price dependency rather than external competitiveness. We make use of a new commodity price index devised using the most updated trade weights at our disposal for a group of more than forty commodities across a panel of seventy-two countries.

The chapter is organized as follows: Sections 3.1 and 3.2 underline the contributions and motivations of the chapter. In Subsections 3.2.1 and 3.2.2 we find an empirical literature review together with the theoretical framework we borrowed from Cashin, Céspedes, and Sahay (2004). Section 3.3 presents our database and focuses on the variables, explaining how we sourced our exchange rate series and how the commodity price index was built. Section 3.4 reviews the statistical time properties of the series and shows the results of the panel unit root and cointegration tests. Section 3.5 presents the dynamic OLS panel estimates for the whole database; Subsection 3.5.1 contains a robustness check using an alternative definition of the exchange rate, while Subsection 3.5.2 a preliminary group analysis. Section 3.6 explores the effects of accounting for cross sectional dependence as well as coefficient heterogeneity in the whole dataset (Subsection 3.6.1) and in the country groups (Subsection 3.6.2). Section 3.7 shows the panel Granger causality tests and the weak exogeneity tests. Section 3.8 concludes the chapter.

3.2 Motivations

3.2.1 Empirical literature review and motivations

Panel data applications have been fundamental in developing a more accurate analysis of the long run evolution and short term convergence of the real exchange rate to its equilibrium. However, determinants of the exchange rate across countries have generally been subject to the not always advisable restriction of a common long run behavior. Past studies have stated that the real exchange rate could be approximated, when considering its slower than expected convergence rate to its equilibrium,[1] by a non linear univariate process that mimics the existence of menu costs (Michael, Nobay, and Peel (1997), Obstfeld and Taylor (1997b), Taylor, Peel, and Sarno (2001a), Imbs, Mumtaz, and Ravn (2003)). However, up until that point little had been said about alternative sources of non linearity and, consequentially, alternative equilibrium relationships, in the real exchange rate adjustment behavior.

In the 2000s, evidence of a long run relationship between real exchange rates and commodity prices had found consistent, albeit country case limited, confirmation in the seminal works of Chen and Rogoff (2003b), who first tested such relationship for

[1] See, needless to say, Rogoff (1996) for an overview of the main puzzles of empirical international economics, which would include this issue.

a group of five advanced countries, and Cashin, Céspedes, and Sahay (2004), who switched the attention to commodity producing countries. Recent evidence of the relationship between commodity prices and real exchange rates can be found in Ricci, Milesi-Ferretti, and Lee (2013), as the authors tested for cointegration and found a positive and statistically significant long run elasticity in a panel of forty-eight industrialized and emerging countries, where a commodity terms of trade was devised and calculated at yearly frequency. More recently Bodart, Candelon, and Carpantier (2015b) and Bodart, Candelon, and Carpantier (2012) used non-stationary panel data techniques in a large monthly panel off sixty-eight countries, and found evidence of a threshold export weight of around twenty percent, above which the relationship between exchange rate and commodity prices would exist.[2]

While deciding whether or not to include given countries in a sample generally depends on homogeneous criteria, like GDP per capita grouping, export share and emerging/non emerging country definitions, it is potentially interesting to take advantage of modern averaging and pooling/averaging panel data techniques for large T panels to uncover more of the relationship between commodities and exchange rates by relaxing some of the homogeneity assumptions on the structure of the long run relationship between such variables. Our analysis starts by reporting the following first step, cointegrating equation specification, which pretty much nests every cross sectional and panel data attempt at describing real exchange rate ($y_{i,t}$) convergence:

$$y_{i,t} = \alpha_1 + \delta_{1(i)} X_{i,t} + f_i + f_t + u_{i,t} \tag{3.1}$$

which can be easily reparametrized and leads to the following error correction model:[3]

$$\Delta y_{i,t} = \alpha_2 + \delta_{2(i)} ET_{(i),t-1} + \delta_{3(i)} \Delta y_{i,t-1} + \delta_{4(i)} \Delta X_{i,t-1} + \varepsilon_{i,t} \tag{3.2}$$

where $ET_{(i),t-1} = y_{i,t-1} - \alpha_1 - \delta_{1(i)} X_{i,t-1} - f_i - f_t$. Equation (3.2), taking into account the existence of unconditional convergence and the effect of volatility on exchange rate fluctuations, or in other words absence of conditional adjustment ($ET_{(i),t-1} = 0$), can

[2] For an overview of past literature results, see Table 3.2.1.

[3] For simplicity we offer a one lag rendition of the model. Higher dynamics can be specified.

be dynamically re-specified starting over from (3.1) as:

$$\Delta y_{i,t} = \alpha_3 + (\beta_{(i)} - 1)y_{i,t-1} + \delta_{5(i)}\Delta X_{i,t} + f_i + f_t + e_{i,t} \tag{3.3}$$

where f_i and f_t would represent individual and regional specific time effects dummies, the coefficient $\delta_{2(i)}$ measures conditional convergence or adjustment speed and $(\beta - 1)$ can be thought as the real exchange rate attractor term, which is assumed to be negative by theory and captures the concept of unconditional convergence. $X_{i,t}$ represents a vector of controls based on behavioral definition of the exchange rate, such as a measure of the degree of trade openness, the net foreign asset position, the ratio of government consumption over GDP, a measure of price inflation, and the terms of trade. Estimation of equations (3.1) and (3.3) has been carried out across empirical applications with various techniques, and has found empirical validity in most of its specifications if not all, leaving some window of ambiguity over the magnitude and direction of some of them. As we already said, much attention has been put in the past in deriving a terms of trade variable and a variety of price indexes, which would normally enter the exchange rate equation as fundamental controls. Bodart, Candelon, and Carpantier (2015b) and Bodart, Candelon, and Carpantier (2012) suggest an interesting change in such set up allowing for the price of a selection of leading exported commodities to enter the exchange rate equilibrium equation.

This chapter[4] borrows from them such idea an tries to capture the relationship between commodity prices and real exchange rates by updating the commodity price index variable created by Cashin, Céspedes, and Sahay (2004) with a set of newly updated price weights, which we sourced from the same authors.

Procedures relying on a combination of fixed effects estimators and instrumental variables, normally employed in “small N, large T” panels, like the Arellano and Bond GMM (generalized method of moments) estimator, would normally restrict all slope coefficients to be the same, assume homogeneous time effects and above all, treat errors as cross-sectionally independent. The last assumptions, especially in a macroeconomic framework where prices dynamics are analyzed, would appear to be extremely restric-

[4] And, as you shall see, the following.

tive.[5] Furthermore, in past applications, the use of dynamic lag methods based on an error correction representation solves another sensible issue, which was first found to be related to growth studies and the required time span selection. As Islam (1995) remarks, the correct choice of the time span is critical when it comes to avoid short term disturbances to acquire an out of scale magnitude in a brief time span, thus making empirical studies on convergence less reliable. Even though being consistent and easily applicable to all panel estimators, the common restriction of averaging up to five years long non-overlapping periods so persistent in past literature can be avoided using estimators from the mean group family. To our knowledge, literature related to such approach has been relatively scarce up until now. As a methodological contribution and in order to tackle the issues of the specific country time effects and the cross sectional dependence while testing for the degree of homogeneity of long run elasticities, we begin our analysis using the mean group DOLS estimator by Pedroni (2001), and later resort to three estimators from the so called pooled mean group family of estimators,[6] and employ preliminary unit root tests and cointegration tests which take possible cross sectional correlation into account.

3.2.2 Theoretical framework

This chapter employs a definition of real exchange rate commonly used in literature, based on the regressors chosen to account for its behavioral equilibrium definition by Macdonald (1998) and Macdonald (2000). The benchmark equilibrium condition which mostly characterizes the literature on the exchange rates is the so called purchasing power parity (PPP). Using for simplicity home and away price notation P_1 and P_2^*, the real exchange rate can be though as the relationship between two price baskets in levels:

$$P_1 = \varepsilon P_2^* \tag{3.4}$$

where ε represents the nominal bilateral exchange rate. Following the absolute version of the purchasing power parity, as the nominal exchange rates should be equal to one,

[5] We will carefully explain that by addressing why a subcript i for the coefficient has been used in equations (3.1) and (3.3).

[6] For a different view on this kind of averaged estimators, have a look at Eberhardt (2012) or Eberhardt and Teal (2011), which mainly focus on growth and development applications.

Table 3.1: Literature's results

ARTICLE	SELECTION, TIME, RESULTS
Amano and Van Norden (1995)	Small, exporting commodity country
Canada	Annual 1973 - 1992
Commodity Terms of Trade	0.8 (non-energy only)
Chen and Rogoff (2003)	Assumption of small economy exogeneity
Australia, Canada, new Zealand	Quarterly, 1984 - 2001
Real Commodity Price index	0.51 to 2.16
Cashin et al. (2004)	50% export earnings (non-fuel)
53 developing countries, 5 advanced	Annual, 1980-2000
Real Commodity Price index	0.1 to 2
Ricci et al. (2013)	IMF total availability
48 industrial and emerging countries	Annual, 1980 - 2004
Commodity Terms of trade	0.46 to 0.76
Bodart et al. (2012)	20% of total export (single commodity)
68 commodity countries	Monthly, 1980 - 2009
Single Real Commodity Prices	0.16 to 0.63
Bodart et al. (2015)	20% of total export
33 developing countries	Annual, 1980 - 2012
Real Commodity Price index	0,102 to 0,372
Coudert et al. (2015)	From Cashin et al. (2004), with additions
69 commodity exporting countries	Annual 1980 - 2012
Real Commodity Price index	0.202 linear , 0.0724 non-linear

then the prices of the same (basket of) goods across any given country should be the same, once a numeraire has granted conversion to the same unit value:

$$P_1 = P_2^* \tag{3.5}$$

or, equivalently

$$\frac{P_1}{\varepsilon P_2^*} = \frac{P_1}{P_2^*} \tag{3.6}$$

where the ratios above represent two equivalent definitions of the purchasing power parity. In principle, we would like to work with both constant values and level variables. However, given the scope of our paper and considering our prices are time indexed, what we actually tested for is the relative version of the purchasing power parity:

$$\frac{\Delta p_{1,t}}{\Delta p_{2,t}^*} = z_t \longrightarrow k \tag{3.7}$$

Such equality, in absolute or relative terms, would constantly hold for any time span. However, as the Penn effect states, the price of a given product will permanently stay higher whenever an economy is characterized by a high(er) income, relative to other countries. In the shade of such empirical disequilibrium, one of the necessary controls entering real exchange rate related specifications is a proxy for productivity differentials that affect the tradable sector and the relationship between prices and wages in it, thus determining the difference. Given the marginal product of labor in country one and two for the non tradable ($MPL_{NT,1}; MPL_{NT,2}$) and tradable ($MPL_{T,1}; MPL_{T,2}$) sectors, and considering prices of non tradables ($p_{NT,1}; p_{NT,2}$) and the international price of tradables (p_T), and assuming the labor market is flexible enough to ensure mobility across markets so that wages equalize, if productivity in the non tradable sectors is the same in countries one and two:

$$MPL_{NT,1}; MPL_{NT,2} \tag{3.8}$$

but the tradable sector in country two suffers, for any given reason, a productivity shock such that:

$$MPL_{T,1} > MPL_{T,2} \tag{3.9}$$

then, for the following equilibrium to hold:

$$\begin{aligned} w_1 &= p_{NT,1} * MPL_{NT,1} = p_T * MPL_{T,1} \\ w_2 &= p_{NT,2} * MPL_{NT,2} = p_T * MPL_{T,2} \end{aligned} \tag{3.10}$$

we will necessarily need to have $p_{NT,1} > p_{NT,2}$.

That is, prices in the non tradable sector will adjust as the marginal product increases to equalize wages across sectors. This will create, under sufficient labor flexibility, a constant price gap in the aggregate price index $P = p_{NT}^{\alpha} * p_T^{1-\alpha}$ that will lead to systematic deviations in terms of real prices and to a "failure of the PPP". In others words, to higher productivity we would expect to see a higher (more appreciated) real exchange rate. In our framework and following past literature, productivity differentials are controlled by proxying them with the logarithm of real GDP per capita at current price levels.[7]

Gains and losses in terms of international competitiveness are usually captured by a terms of trade variable (TOT). The relationship between export and import prices is linked to the performance gap in productivity of non tradable sectors. In particular, Chen and Rogoff (2003b) finds that the relative (international) price of nontradables appears to be correlated to relative export prices, and as such the non tradable price index would increase with export prices the same way it would if the marginal product of labour in the tradable sector was higher, thus following the wage/prices equilibrium dictated by the Balassa-Samuelson effect.

As for the transfer problem (Keynes 1929, Ohlin 1929), a net foreign assets variable (NFA) controls for excessive liabilities that might cause a long run deterioration of the terms of trade and an appreciation of the exchange rate through changes in the composition of consumption preferences across sectors in the country that receives the transfer. As not all the transfer is spent in export goods from the liable country, additional demand in the non tradable sector will draw labour out of the tradables

[7] A choice we share in common with Coudert, Couharde, and Mignon (2015), for a recent example.

sector, making prices and relative wages and export prices increase with respect to import prices. As a result, in the foreign creditor country, some part of the export will be redirected to the nontradables, and in the home liable country some part of the nontradable goods will be redirected into exports. This will imply a deterioration of the terms of trade for the home countries and an appreciation of the real exchange rate. This position, held by Keynes, is usually counterbalanced by standard inter-temporal models, which assume that countries which held external liabilities will want to depreciate the exchange rate in order to get to a trade surplus which will allow for repayment of the debt in the long run.

We resort to the ratio of Government expenditure to GDP (GOV) to serve as a proxy for financial stability. Contrary to any indicator variable present in literature which only allows for a sharp regime change, from a less to a more open one, we hope this simple measure could adequately capture gradual shifts in financial risk and the propensity of a country to shift upwards its default risk by increasing deficit spending.

As a last theoretical point, we would need to explain how the real exchange rate relates to real commodity prices. Following Cashin, Céspedes, and Sahay (2004), and after some math:

$$\frac{P_1}{\varepsilon P_2^*} = \left(\frac{MPL_{EXP,1}}{MPL_{IMP,2}} \frac{MPL_{NT,2}}{MPL_{NT,1}} \frac{P_{EXP,2}}{P_{IMP,2}} \right)^{\lambda} \tag{3.11}$$

which would equal:

$$\ln\left(\frac{P_1}{\varepsilon P_2^*}\right) = \lambda \ln\left(\frac{MPL_{EXP,1}}{MPL_{IMP,2}} \frac{MPL_{NT,2}}{MPL_{NT,1}} \right) + \lambda \ln\left(\frac{P_{EXP,2}}{P_{IMP,2}}\right) \tag{3.12}$$

as productivity differential are captured by the constant term, the baseline specification will thus be:

$$\ln\left(\frac{P_1}{\varepsilon P_2^*}\right) = k + \lambda \ln\left(\frac{P_{EXP,2}}{P_{IMP,2}}\right) \tag{3.13}$$

introducing terms of trade, NFA and export to GDP the final specification thus amounts to:

$$REER = \ln\left(\frac{P_1}{\varepsilon P_2^*}\right) = k + \lambda \ln\left(\frac{P_{EXP,2}}{P_{IMP,2}}\right) + \gamma X + \varepsilon \tag{3.14}$$

with $X = NFA, GOV$ and/or any other additional covariate accounted for in literature.

Equation (4.14) simply states that the real exchange rate ($REER$) moves along with any (commodity) price basket evaluated at foreign (intermediate) prices. International export prices, in this theoretical set up, where we do focus on commodity exporters, represent an exogenous source of fluctuation for the exchange rate, and affect prices in the non tradable markets as well as wages across all sectors given they are perfectly flexible. Note that in a framework making use of a terms of trade variable, such relationship would imply in real terms that:

$$TOT = \frac{P_{EXP,2}}{P_{IMP,2}} \approxeq \frac{P_{EXP,2}}{MUV} \tag{3.15}$$

so that the terms of trade variable (TOT) would stand for external competitiveness, with $P_{EXP,2}$ being defined as the nominal export commodity price index while foreign intermediate prices of imported goods ($P_{IMP,2}$) would be proxyed by the Manufacture Unit Value index (MUV), a measure close to the consumer price index of the United States, which acts as a deflator. A commodity terms of trade has been indeed been employed in past literature as much as a standard prices to search for the optimal long run covariates capable of describing the exchange rate, and abstracting from monetary based models of determination as those recently examined in Sarno and Schmeling (2014).

3.3 Database and variables

We follow Cashin, Céspedes, and Sahay (2004) exercise to construct an export-commodity index of forty commodities using the most updated trade weights at our disposal. The index calculation was thus obtained by summing across the MUV-deflated weighted average of each individual commodity price by its own export weight. As an additional contribution to this area of study, we calculated the index using newer weights averaged in the 1999-2004 period, as they were made available by Cashin, Céspedes, and Sahay (2004), and have never been put to use. Our Commodity (Export) Price Index was as such calculated as:

$$NCOMP_{i,t} = \exp\left[\sum_{k=1}^{K} Weight_{i,k} * \ln(price_k)\right]_t \tag{3.16}$$

and:

$$COMM_{i,t} = \frac{\exp\left[\sum_{k=1}^{K} Weight_{i,k} * \ln(price_k)\right]_t}{MUV_t} \tag{3.17}$$

where $NCOMP_{i,t}$ represents the nominal commodity price index at time t within country i, $COMM_{k,t}$ represents the real commodity price index at time t within country i, $Weight_k$ is the average 1999-2004 average export weight of commodity k on the total of exports of goods of country i, $price_k$ is the international price of exported commodity k, and finally MUV_t represents the Manufacture Unit value index at time t used to deflate the Nominal prices $NCOMP_{i,t}$.

As for the real effective exchange rates we employed, our solutions were the CPI based real effective exchange rate from the international financial statistics database (IFS) of the IMF and, as a main benchmark, the recently devised and continuosly updated real exchange rate by Darvas (2012). In particular, following the $REER$ definition by Darvas (2012), we consider:

$$REER_{i,t} = \frac{NEER_{i,t} * CPI_{i,t}}{CPI_t^F} \tag{3.18}$$

where $REER_{i,t}$ is the real exchange rate of country i at time t evaluated against a basket of foreign currencies, $CPI_{i,t}$ is the consumer price index of country i at time t, and $NEER_{i,t}$ is the nominal exchange rate of country i in period t and finally CPI_t^F represents a weighted average of all the CPI's from trade partners for every period t. Both the consumer price index and the nominal effective exchange rate are in turn the geometrical averages of the country i's trade partners' CPI indexes and country i's bilateral exchange rates with trade partners respectevely:

$$NEER_{i,t} = \prod_{f=1}^{N} NER_{f,t}^{(w_i)} \tag{3.19}$$

$$CPI_t^F = \prod_{f=1}^{N} CPI_{f,t}^{(w_i)} \tag{3.20}$$

where $NER_{f,t}^{(w_i)}$ is the nominal bilateral exchange rate of country i with respect to foreign partner f, weighted by the country weight w_i, and $CPI_{f,t}^{(w_i)}$ represents the consumer price index of trade partner f at time t weighted by home country weight w_i. Once again, $\sum_{i=1}^{N} w_i = 1$, and the weights are time invariant, as before.

The variables we have taken from the World Bank's World Development Indicators (WB WDI) database are the net foreign asset position (NFA), real per Capita GDP (GDP) and finally the share of government spending over GDP (GOV). Many, if not the majority of the literature we reviewed, have up until now consistently chosen years between 1980 and 1995 as the initial sampling period. Historical reasons accounted for are the beginning of the globalization era, a more commodity price insulated common global structure, as well as availability of high quality data. Given current data availability, our database spans a set of balanced observation ranging monthly from 1995 to 2012.

3.4 Unit root tests and cointegration

All the variables we tested were found to be integrated of order one. As we checked for levels and first differences to check for stationarity, we followed the sequential algorithmic procedure as described in Dickey and Pantula (2002). We employed two tests, the Maddala and Wu (1999) first generation unit root test and the Pesaran (2007) CIPS test, a second generation test which would account for possible cross-sectional correlation in our specification. We tested for a series of alternative specifications which included up to six lags and both a trending and non trending alternative. In this section, we report the results for the whole panel in Table 4.1.

In Table 4.2, we also report the results from the panel cointegration tests. As our analysis focuses on long run elasticities, we employ the Westerlund (2007) cointegration tests borrowing the methodology from Blackburne and Frank (2007), which allows us to test for cointegration through an error correction specification, calculating tests that do not only present an homogeneous no cointegration null (P_t and P_a), but also an alternative couple of heterogeneous tests whose alternative hypothesis would imply that only some of the sections are cointegrated (G_t and G_a). To account for ex-ante eventual cross sectional correlation, we follow the author and *bootstrap the p-values*, repeating the test one hundred times. We repeat the analysis five times following a top down identification strategy, in order to infer on the existence of a possible relationship between the real

Table 3.2: Unit root tests

$Variable$	$Maddala\ and\ Wu$ (1999)	$P-value$	$Pesaran$ (2007)	$P-value$
$REER$	156.187	0.166	-0.097	0.461
$COMM$	56.673	1.000	0.568	0.715
GDP	129.817	0.720	8.091	1.000
NFA	132.088	0.671	-1.387	0.083
GOV	139.166	0.504	0.420	0.663

Unit root tests in levels, null hypothesis: the variable has a unit root. Up to six lags tested. The alternative for the Pesaran test includes a trend.

exchange rates and commodities. Overall results for the two homogeneous alternative versions of the tests, P_t and P_a, reject the unit root null anywhere in the panel, and show widespread cointegration in it. With an heterogeneous alternative, an exception is visible for the just slightly higher result in the G_a group test for the commodities, which would imply that some of the cross sections might not be cointegrated as we expect, and thus further justify the use of the Pedroni's Mean Group DOLS estimation the following section.

3.5 Whole panel

Tables 3.6 and 3.7 present the full model for the whole database, showing long run elasticities as well as the error correction terms estimated by a first step superconsistent ordinary least square estimation. The alternative specifications follow the multistep cointegration analysis and were ordered with the intention to test for the strength of the relationship between the exchange rate and our commodity index. Overall results do indicate a consistently positive and significant elasticity of the real exchange rate with respect to export commodity prices.

Table 3.3: Cointegration tests

$REER$	G_t	$P - value$	G_a	$P - value$
$COMM, GDP, NFA, GOV$	-2.874	0.020	-15.655	0.030
$COMM, GDP, NFA$	-2.613	0.010	-13.034	0.010
$COMM, GDP$	-2.322	0.040	-10.661	0.040
$COMM$	-2.119	0.040	-8.567	0.070
$REER$	P_t	$Pvalue$	P_a	$Pvalue$
$COMM, GDP, NFA, GOV$	-23.698	0.040	-13.928	0.010
$COMM, GDP, NFA$	23.094	0.000	-12.706	0.000
$COMM, GDP$	-20.995	0.010	-10.507	0.020
$COMM$	-18.486	0.010	-8.291	0.010

Cointegration tests. Robust P-values are bootsptrapped one hundred times. Constant included. Pt and Pa tests both have a homogeneous null of non stationarity. Gt and Ga have a heterogeneous null of non-stationarity in variance: at least one section shows cointegration.

Table 3.4: Pedroni mean group DOLS estimates

COMM	0.064	0.069	0.034	0.059
	(6.202)	(5.904)	(3.888)	(1.083)
GDP	0.217	0.271	0.256	
	(10.810)	(11.560)	(16.450)	
NFA	0.111	-0.003		
	(-3.920)	(-3.922)		
GOV	0.267			
	(13.130)			
ECT	-0.017	-0.017	-0.016	-0.018
	(-13.230)	(-13.240)	(-12.980)	(-14.080)
Observations	13242	13242	13242	13253
Countries	70	70	70	70
Time demeaning	YES	YES	YES	YES
Max Lags	6	6	6	6
Trend	YES	YES	YES	YES

Pseudo t-ratio statistics in parentheses. ECT stands for error correction term, the speed of adjustment calculated from a linear error correction specification.

The net foreign assets elasticity for the whole sample is around 0.111, which would mean that a 10% increase in the net foreign asset would be associated with a long run appreciation of the real exchange rate of around 1%. Such value appear to be around a half of the value that Ricci, Milesi-Ferretti, and Lee (2013) have estimated. A catastrophic deterioration of 50% would thus imply an exchange rate deterioration of around 5 percentage points.

In line with Coudert, Couharde, and Mignon (2015) and Ricci, Milesi-Ferretti, and Lee (2013), a 10% increase in relative productivity, as proxyed by the GDP, would be associated with a long run appreciation of the real exchange rate of around 10%. Such value is perhaps in line with the literature and stays consistent throughout our estimates with values ranging from 0.217 to 0.256.

Government consumption is estimated to appreciate the exchange rate as well in the log run, as an increase in spending would be associated with an appreciation we estimate to be much lower than in the seminal papers of De Gregorio and Wolf (1994) and De Gregorio, Giovannini, and Wolf (1994), which showed a 1.5 to 3% appreciation effect and to the 3% appreciation of Ricci, Milesi-Ferretti, and Lee (2013). Such effect arguably depends on the composition of the panel, whereas less advanced and financially burdened countries are coupled with more advanced and generally financially constrained ones.

The commodity terms of trade presents a consistent and almost always significant elasticity ranging from 0.034 to 0.069. A 10% increase in our commodity index would imply an appreciation ranging between 0.34 to almost 0.7%. Such values are undoubtedly lower than what has been previously estimated by Ricci, Milesi-Ferretti, and Lee (2013) and by Cashin, Céspedes, and Sahay (2004), which would detect an average appreciation of 4.2% after a 10% increase in commodity prices, with an average elasticity of around 0.42. With updated weights and allowing for a higher degree of cross sectional heterogeneity, our estimated coefficient appears to be, indeed, much lower. Arguably, as economies become more integrated and future contracts substitute spot purchases, our results show how the ten year long period which our analysis adds to the Cashin, Céspedes, and Sahay (2004) timespan has seen world economies becoming more resilient to primary goods price shocks.

Table 3.5: Pedroni mean group DOLS estimates, IFS REER

COMM	0.015	0.035	-0.219	0.027
	(2.298)	(3.406)	(-2.152)	(-3.754)
GDP	0.197	0.316	0.380	
	(9.644)	(10.490)	(16.440)	
NFA	-0.120	-0.101		
	(-6.284)	(-5.444)		
GOV	0.024			
	(13.540)			
ET	-0.016	-0.016	-0.016	-0.015
	(-12.980)	(-12.890)	(-12.900)	(-12.880)
Observations	13289	13289	13289	13300
Countries	70	70	70	70
Time demeaning	YES	YES	YES	YES
Max Lags	6	6	6	6
Trend	YES	YES	YES	YES

Pseudo t-ratio in parentheses. Lag selection carried out through AIC.

Finally, our implied speed of convergence does suggest a much slower speed of adjustment to equilibrium, suggesting a sluggish persistence of price shocks once they have entered the economy and affected its relative prices.

3.5.1 Robustness check

In order to add to the robustness of our results, we employ an alternative definition of real exchange rate, sourced directly from the International Monetary Fund, visible in Table 3.5. Although with some differences, we see how commodity prices still present lower-than-previous-literature elasticities, amounting to around half the value of the main estimates and providing a clear confirmation of the weakening of the degree of dependency of the exchange rate to commodity price variations, while the speed of adjustment stays slow and comparable to the main estimates.

3.5.2 Group analysis

We now proceed to discuss the group results, with a focus on advanced, European and emerging economies under the light of past literature grouping. As a matter of fact, once distinct groups of countries are taken into account, our results shows some degree of consistency with past literature (again, Bodart, Candelon, and Carpantier (2012), Ricci, Milesi-Ferretti, and Lee (2013), Coudert, Couharde, and Mignon (2015), Cashin, Céspedes, and Sahay (2004)), effectively underlining strong differences between regional groups and commodity thresholds in an aggregated panel study.[8]

In Tables 3.6 and 3.7 we have carried out estimates for an available sample of Cashin, Céspedes, and Sahay (2004) countries (Column I), Ricci, Milesi-Ferretti, and Lee (2013) selection of advanced countries (Column II), Ricci, Milesi-Ferretti, and Lee (2013) advanced non EU countries (Column III), Ricci, Milesi-Ferretti, and Lee (2013) advanced EU countries (Column IV), Ricci, Milesi-Ferretti, and Lee (2013) nearly industrialized and emerging economies (Column V), an augmented sample of all non advanced economies at our disposal (Column VI), an augmented sample of all advanced economies at our disposal (Column VII), a group of major energy exporters in accordance with the

[8] *Group I* contains the following countries: Australia, Bolivia, Brazil, Central African Republic, Canada, Chile, Cote d'Ivoire, Cameroon, Colombia, Costa Rica, Ghana, Morocco, Mexico, Malaysia, Nicaragua, Norway, New Zealand, Pakistan, Philippines, Paraguay, Tunisia, Uganda, Uruguay and South Africa. *Group II* contains Australia, Austria, Belgium, Canada, Switzerland, Germany, Denmark, Spain, Finland, France, United Kingdom, Greece, Ireland, Italy, Japan, Netherlands, Norway, New Zealand, Portugal, Sweden and United States. *Group III* contains Australia, Canada, Switzerland, Denmark, United Kingdom, Japan, Norway, New Zealand, Sweden and United States. *Group IV* contains Austria, Belgium, Germany, Spain, Finland, France, Greece, Ireland, Italy, Netherlands and Portugal. *Group V* contains Brazil, Chile, China, Colombia, Czech Republic, Hungary, Morocco, Mexico, Malaysia, Pakistan, Philippines, Poland, Russia, Singapore, Slovak Republic, Venezuela and South Africa. *Group VI* contains: Bulgaria, Bahrain, Belize, Bolivia, Brazil, Central African Republic, Chile, China, Cote d'Ivoire, Cameroon, Colombia, Costa Rica, Cyprus, Czech Republic, Dominican Repulbic, Algeria, Gabon, Georgia, Ghana, Gambia, Guyana, Hungary, Iran, Latvia, Morocco, Moldova, Mexico, Macedonia, Malta, Malaysia, Nigeria, Nicaragua, Pakistan, Philippines, Poland, Paraguay, Romania, Russia, Saudi Arabia, Singapore, Slovak Republic, Tunisia, Uganda, Ukraine, Uruguay, Venezuela and South Africa. *Group VII* contains: Australia, Austria, Belgium, Canada, Switzerland, Germany, Denmark, Spain, Finland, France, United Kingdom, Greece, Ireland, Iceland, Israel, Italy, Japan, Netherlands, Norway, New Zealand, Portugal, Sweden and United States. *Group VIII* contains: Bahrain, Bolivia, Cote d'Ivoire, Cameroon, Colombia, Algeria, Gabon, Ghana, Guyana, Iran, Mexico, Nigeria, Norway, Russia and Saudi Arabia. *Group IX* contains: Australia, Austria, Belgium, Bulgaria, Belize, Brazil, Central African Republic, Canada, Switzerland, Chile, China, Costa Rica, Cyprus, Czech Republic, Germany, Denmark, Dominican Republic, Spain, Finland, France, United Kingdom, Georgia, Gambia, Greece, Hungary, Ireland, Iceland, Israel, Italy, Japan,Latvia, Morocco, Moldava, Macedonia, Malta, Malaysia, Nicaragua, Netherlands, New Zealand, Pakistan, Philippines, Poland, Portugal, Paraguay, Romania, Singapore, Slovak Republic, Sweden, Tunisia, Uganda, Ukraine, Uruguay, United States and South Africa.

Table 3.6: Pedroni mean group DOLS estimates, country selection I

	(I)	*(II)*	*(III)*	*(IV)*	*(V)*
COMM	0.061	0.078	-0.070	0.057	0.116
	(3.692)	(4.185)	(0.516)	(2.778)	(2.782)
GDP	0.671	0.126	0.978	0.209	1.026
	(10.800)	(1.118)	(1.640)	(3.981)	(12.070)
NFA	-0.008	-0.243	-0.307	-0.739	0.134
	(4.151)	(-6.900)	(-6.474)	(-8.157)	(-2.048)
GOV	0.042	-0.132	-0.067	0.009	0.501
	(7.914)	(1.835)	(4.546)	(3.940)	(13.030)
Observations	4513	3990	1710	2090	3230
Countries	24	21	9	11	17
Time demeaning	YES	YES	YES	YES	YES
Max lags	6	6	6	6	6
Trends	YES	YES	YES	YES	YES

Pseudo t-ratio in parentheses. Lag selection carried out through AIC.

Table 3.7: Pedroni mean group DOLS estimates, country selection II

	(VI)	(VII)	$(VIII)$	(IX)
$COMM$	0.065	0.057	0.215	0.040
	(4.921)	(4.156)	(5.213)	(0.963)
GDP	0.580	-0.822	-0.217	0.533
	(11.820)	(0.135)	(5.801)	(12.880)
NFA	0.162	-0.129	0.009	0.018
	(4.066)	(-5.210)	(0.837)	(-2.411)
GOV	0.415	-0.323	0.012	0.310
	(14.500)	(1.469)	(2.881)	(12.600)
Observations	9073	4359	3040	10202
Countries	48	23	16	54
Time Demeaning	YES	YES	YES	YES
Max lags	6	6	6	6
Trends	YES	YES	YES	YES

Pseudo t-ratio in parentheses. Lag selection carried out through AIC.

international monetary fund's world economic commodity database (WECD) and mimicking Bodart, Candelon, and Carpantier (2012) with a threshold export share of 20% (Column VIII), a group of major non energy exporters according to the IMF WECD mimicking Bodart, Candelon, and Carpantier (2012) with a threshold export share of 20% (Column IX).

As we look at Column VI and VIII and focus once more on commodity prices, evidence shows how our updated estimates close up to literature only when non advanced economies and energy exporters with an export share larger than 20% are considered. The estimates would imply that a 10% increase in the commodity index would cause a long run appreciation of 6.5% in emerging economies, as well as a 2.15% appreciation in energy export dependant countries. Such results are in contrast with past literature results, as a commodity term of trade shock should affect the exchange rate in larger magnitude in advanced economies, and can be explained in terms of export dependency and sensitivity to export price changes rather than competitive shifts in export prices relative to import prices. The difference in terms of elasticity is clearly much less visible, although to a lesser degree given differences in sample size, in columns I and II, where we stick to countries which were included by Ricci, Milesi-Ferretti, and Lee (2013).

In Column I, the group sample of Cashin, Céspedes, and Sahay (2004) countries brings an estimate of 0.061, that is almost half less than the lower elasticity the authors estimated in their application. The extended sample of non-energy exporting markets with a threshold of 20%, mimicking Bodart, Candelon, and Carpantier (2012) discretional selection, closely follows such results with a long run elasticity of 0.040. Perhaps the only result which falls in line with previous literature can be seen in Column VI. Countries which are considered non advanced by the IMF, may them be net importers or exporters of commodities, show an estimated elasticity of 0.651, implying that a 10% increase in the commodity price index would have an appreciation effect of around 6.5% on the exchange rate. All additional covariates have a fairly acceptable magnitude and positive sign, with a 1% increase in goverment spending associated with an appreciation of 0.58%[9] and a net foreign asset elasticity of 0.162, implying a 1.6% appreciation of

[9] Much lower than what Ricci, Milesi-Ferretti, and Lee (2013) and De Gregorio, Giovannini, and Wolf (1994) found out about emerging economies, with GDP elasticities ranging from 1,5 to almost 4, implying a strong, more than one to one relationship.

the rate following an increase of external assets of 1%.[10]

3.6 Exchange policy spillovers and market integration

3.6.1 Whole panel results

The results we have seen up until now focus on the idea that, ceteris paribus, markets in commodity dependent countries have remained partially non integrated and as such represent a direct driver of currency (mis)alignment. However, such way of thinking about countries' policies as univocal choices is not necessarily correct. In reality, it is impossible to think about neighboring agents not being influenced by common time varying factors, albeit being unobserved. In particular, time varying factor (t-indexed ones) would qualify as common shocks or proximity spillover effects, which in our environment affect productivity of neighboring countries, leading to more or less common choices in terms of trade openness and exchange rate policy across countries. Biasedness of the estimates we have seen until now is then conditioned by the strength and number of common unobserved factors.

In this chapter, we tackle the issue related to whether or not capturing the heterogeneity[11] in the discretional and automatic rules which decide the intensity of the relationship between the exchange rate and the commodity prices might not be enough to account for an unbiased estimation if at least some form of common behavior has not been accounted for. If we accept the fact that at least a single common but weak heterogeneous-effect shock might have conditioned the equilibrium between the exchange rate and the commodity prices, than we are stating that at least a small bias has been present in past literature estimates. Let us assume the following structure for the intended relationship between the exchange rate and its fundamentals:

$$y_{i,t} = \beta_i x_{i,t} + u_{i,t} \tag{3.21}$$

[10] This appears to be a higher estimate compared to past theory using the ratio of NFA to trade but somehow in line with previous NFA to GDP estimates (see Lane and Milesi - Ferretti (2004)) and with the idea that emerging countries would face sharper movements of their exchange rate due to the tighter borrowing constraints they would be forced to face. Furthermore, as the ratio of average exports plus imports to GDP is somewhat around 25%, our estimated 1,61 in column VI would equal 0,04 if NFA was normalized to trade, a result pretty much in line with previous estimates.

[11] Rather than nonlinearity, to which we dedicate Chapter 4.

$$u_{i,t} = \lambda_i f_t + \psi_i + \varepsilon_{i,t} \tag{3.22}$$

where f_t represents the common time varying factor with heterogeneous (cross sectional) loadings[12], while ψ_i represent a sequence of time invariant shocks, affecting univocally all cross sections. If such structure for the innovation $u_{i,t}$ is also common to the regressor:

$$x_{i,t} = \varrho_i f_t + \pi_i g_t + \phi_i + e_{i,t} \tag{3.23}$$

than such set-up, which has been defined cross sectional dependence, implies the existence of a common shock affecting both the exchange rate $y_{i,t}$ and its determinants $x_{i,t}$. This ultimately implies a natural bias which is proportional to the strength and extent of the common shock: this is clearly visible by solving the above equation for the shock and plucking it into the first two equations above. Thus, reparametrizing the model, we get that:

$$y_{i,t} = (\beta_i + \lambda_i \varrho_i^{-1}) x_{i,t} + \psi_i - \lambda_i \varrho_i^{-1} \phi_i + \varepsilon_{i,t} - \lambda_i \varrho_i^{-1} \pi_i g_t - \lambda_i \varrho_i^{-1} e_{i,t} \tag{3.24}$$

it is evident that $(\beta_i + \lambda_i \varrho_i^{-1}) \neq \beta_i$, and the model is not identified (thus biased, depending on the factor loading λ_i) in its coefficients if unobserved factors are not accounted for. In order to account for such natural bias, this section of the chapter integrates the whole panel results of the dynamic OLS estimates with three additional estimators, the Hashem Pesaran and Smith (1995) MG (mean group) estimator, the Pesaran (2006) CCEMG (common correlated effects mean group) estimator, and finally the Eberhardt and Teal (2010) (augmented mean group) AMG estimator. The latter two are capable of taking into account unobserved common factors with heterogeneous loadings by proxying unobservables (and omitted elements of a cointegrating vector in the case of the AMG) of elasticities with cross-section averages of all variables.

Shifting our attention from standard dynamic models to heterogeneous ones allowing for cross section dependence and accounting for omitted variable bias (AMG and

[12] We will assume the number of unobserved common factors to be equal to one during the remainder of the thesis. This will allow for the interpretation of the AMG results, which assume a non identifiable common dynamic factor and also act as a simplifying assumption for the CCMG and MG estimators.

Table 3.8: Whole panel estimates

	DOLS	*MG*	*AMG*	*CCEMG*
COMM	0.064	0.056	0.032	-0.146
	(6.202)	(3.390)	(1.860)	(-0.550)
GDP	0.217	0.119	-0.007	0.065
	(10.810)	(0.550)	(-0.040)	(0.027)
NFA	0.111	0.016	0.088	0.139
	(3.920)	(0.320)	(1.630)	(2.250)
GOV	0.267	0.018	0.016	0.017
	(13.130)	(4.360)	(3.950)	(4.090)
CSD		17.778	6.770	7.127
$p - value$		(0.00)	(0.00)	(0.00)
Observations	14152	14152	14152	14152
Countries	70	70	70	70
Time demeaning	YES	YES	YES	YES
Max Lags	6	-	-	-
Trend	YES	YES	YES	YES

Pseudo t-statistics in parentheses. Common dynamic factor were left unrestricted and are not reported for the AMG estimates. CSD stands for the residual test on cross section dependence.

CCMG) we can see how the overall elasticity of exported goods not only is halved in both the mean group and augmented mean group estimates, but it becomes negative and non-signficant once we take into account the Pesaran CCMG estimator. In all three mean group estimators, post estimation checks was carried out through the Pesaran (2015) weak dependence test. None of the estimators failed to reject the null hypothesis of remaining weak cross sectional dependence. Thus, cross sectional dependence would converge to 0 as the number of cross sectional units increased.

In the case of the AMG, we consider that the presence of a common dynamic factor, a significant one, verifies our hypothesis on market integration: in a more integrated world, where shocks are commonly received in every neighboring market albeit with different loadings, commodity prices impact on the trade and exchange rate policies is mitigated and commodities do not represent the best descriptor for real exchange rate structural trends. Such is the impact on net foreign assets and government spending of common shocks that, once we account for a single weak factor in the CCEMG, we end up getting a totally non significant price index variable. Clearly, the effect of time invariant shocks is difficult to disentangle from that of time varying effects not accounted for in our analysis: estimates from the DOLS and the MG, which do not account for unobservables, are both very similar in shape and size (0.064 against 0.056) and thus we would be inclined to conclude that, although with a new set of time weights and prices, unobserved factor(s) common to all the cross sections represent indeed the real table-turners. However, when compared to the most complete analyses (Cashin, Céspedes, and Sahay (2004), Chen and Rogoff (2003b) and Ricci, Milesi-Ferretti, and Lee (2013) above all), we see how our estimates gave results on average a ten-fold lower than those suggested by the literature, peaking to a twenty-fold with the CCEMG estimates.

Although this might seem logical for a very heterogeneous panel as ours, where emerging, industrialized and commodity dependent countries coexist in the same analysis, more interesting results can be found in the group analysis, where, focusing on specific threshold related characteristics of each group, previous literature results can be confuted.

3.6.2 Group results

Let us now focus our attention on the nine groups we previously estimated through Pedroni's DOLS. In the CCEMG estimator, cross sectional averages augment a regression where model parameters referring to each single group are averaged across the whole panel to get the overall estimator. In the case of the AMG estimator, a first preliminary regression including time dummies is first estimated by OLS in first differences. The resulting coefficient on the time dummies (which the Authors would call *common dynamic process*) are then used as a proxy for unobserved variation in a standard group regression model including a constant, which would capture time-invariant, cross-section specific fixed effects. As a final step, similarly to the basic MG estimation, paramenters from each group are finally averaged to get an overall estimate for the whole panel.

Table 3.9 shows our results for a selected group of emerging countries whose commodity export share amounts to more than twenty percent of their total export, as indicated by the IMF (group IX). In this instance, cross sectional spillovers might be relevant and show signs and magnitude similar to the dynamic estimates, although being in relative terms almost twice as higher, incorporating the magnitude of effect which in the standard estimates was handled by a significant net foreign asset position. In stark contrast to this, but perhaps not surprisingly given the different nature of energy markets, Table 3.10 shows estimates for exporters specialized in energy commodity exports at the 20% threshold (group VIII), which were not found to be significant whatsoever. Once eventual cross sectional spillovers are accounted for by our methodology, the more heterogeneous portfolio of the energy exporting country appears to result in a significant and relatively large in magnitude net foreign asset variable, GDP and government spending stay significant while the commodity index variable, although larger in magnitude, shows a much lower impact compared to what dynamic estimates have shown us, while ranging from insignificant to only marginally significant. Past literature estimates do appear to have overestimated the impact of commodities' terms of trade, especially in energy commodities exporting countries. Our estimates, taking into account possible cross sectional spillover and parameter heterogeneity, presents much lower estimates compared to the standard dynamic OLS estimation which go from halving to reducing to a quarter the magnitude of the effect.

Table 3.9: Non energy commodity exporters, relevant share

	DOLS	*MG*	*AMG*	*CCEMG*
COMM	0.040	0.069	0.030	0.070
	(0.963)	(3.450)	(1.630)	(2.380)
GDP	0.533	0.173	0.187	0.197
	(12.880)	(1.260)	(1.490)	(1.420)
NFA	0.018	0.018	0.007	0.042
	(-2.411)	(0.340)	(0.120)	(0.700)
GOV	0.031	0.017	0.178	0.023
	(12.600)	(3.790)	(3.310)	(3.700)
CDP			0.922	
			(5.050)	
CSD	-	17.605	0.271	2.403
$p - value$	-	(0.000)	(0.786)	(0.016)
Observations	54	10904	10904	10904
Countries	10202	54	54	54
Time demeaning	YES	YES	YES	YES
Max Lags	6	-	-	-
Trend	YES	YES	YES	YES

Pseudo t-statistics in parentheses. CDP indicates the common dinamic process component of the Augmented mean Group estimation. CSD indicates the Cross Sectional Dependence test, with the null alternative of residual cross section dependence. Lag selection carried out through AIC.

Table 3.10: Energy commodity exporters, relevant share

	DOLS	*MG*	*AMG*	*CCEMG*
COMM	0.215	0.012	0.020	0.053
	(5.213)	(0.520)	(0.980)	(1.130)
GDP	-0.217	-0.063	0.314	0.439
	(5.801)	(0.070)	(0.940)	(1.990)
NFA	0.010	0.008	0.284	0.330
	(0.837)	(0.070)	(2.170)	(2.180)
GOV	0.012	0.018	0.022	0.011
	(2.881)	(2.090)	(2.610)	(1.890)
CDP			1.010	
			(2.220)	
CSD	-	3.879	-3.427	-2.611
$p-value$	-	(0.000)	(0.001)	(0.009)
Observations	3040	3248	3248	3248
Countries	16	16	16	16
Time demeaning	YES	YES	YES	YES
Max Lags	6	-	-	-
Trend	YES	YES	YES	YES

Pseudo t-statistics in parentheses. CDP indicates the common dinamic process component of the augmented mean group estimation. CSD indicates the cross sectional dependence test, with the null alternative of residual cross section dependence. Lag selection carried out through AIC.

As a final remark, we show results coming from the Ricci, Milesi-Ferretti, and Lee (2013) selection of emerging and nearly advanced economies (Table 3.11) and the very same selection updated by adding additional countries and time periods made available by the IMF (Table 3.12).

Although with some discrepancies, results from the Ricci, Milesi-Ferretti, and Lee (2013) group of countries stay close to those we obtained with the 20% threshold countries. In any given case, none of the elasticities could get any higher than one percentage point. Results are fairly similar once the last relevant group, made of the Ricci, Milesi-Ferretti, and Lee (2013) countries and augmented with an added number of additional countries sourced from the IMF, comes into play: the commodity variable decreases in magnitude and significance.

Table 3.11: Nearly industrialized and emerging economies

	DOLS	*MG*	*AMG*	*CCMG*
COMM	0.116	0.084	0.030	0.095
	(2.782)	(2.980)	(1.180)	(2.100)
GDP	1.026	0.807	0.758	0.316
	(12.070)	(3.610)	(3.470)	(0.760)
NFA	0.013	0.040	0.120	0.122
	(-2.048)	(0.290)	(1.270)	(0.900)
GOV	0.050	0.040	0.038	0.039
	(13.030)	(4.620)	(4.130)	(4.380)
CDP			0.726	
			(2.870)	
CSD	-	2.744	-6.944	-7.598
$p-value$	-	(0.006)	(0.000)	(0.000)
Observations	3230	3451	3451	3451
Countries	17	17	17	17
Time demeaning	YES	YES	YES	YES
Max Lags	6	-	-	-
Trend	YES	YES	YES	YES

Pseudo t-ratio in parentheses. CDP indicates the common dinamic process component of the augmented mean group estimation. CSD indicates the cross sectional dependence test, with the null alternative of residual cross section dependence. Lag selection carried out through AIC.

Table 3.12: Nearly industrialized and emerging economies, IMF additions

	DOLS	*MG*	*AMG*	*CCEMG*
COMM	0.065	0.060	0.024	0.031
	(4.921)	(3.210)	(1.260)	(1.030)
GDP	0.058	0.177	0.140	0.014
	(11.820)	(0.610)	(0.590)	(0.007)
NFA	0.162	0.106	0.247	0.247
	(4.066)	(1.780)	(3.440)	(2.710)
GOV	0.415	0.020	0.017	0.123
	(14.500)	(4.600)	(4.090)	(2.970)
CDP			0.876	
			(4.250)	
CSD	-	25.740	-2.461	-0.154
$p - value$	-	(0.000)	(0.014)	(0.878)
Observations	9703	9697	9697	9697
Countries	48	48	48	48
Time demeaning	YES	YES	YES	YES
Max Lags	6	-	-	-
Trend	YES	YES	YES	YES

Pseudo t-ratio in parentheses. CDP indicates the common dinamic process component of the augmented mean group estimation. CSD indicates the cross sectional dependence test, with the null alternative of residual cross section dependence. Lag selection carried out through AIC.

3.7 Granger causality and weak exogeneity

To conclude our analysis for this central chapter, we consider a short run Granger causality test based on the work of Dumitrescu and Hurlin (2012) and a short/long run comparative analysis based on the structure of the bilateral error correction specification for the exchange rates and the commodity price indexes. The latter should tell us whether commodity prices and exchange rates are linked by a non-casual relationship and whether or not weak exogeneity of prices with respect to the exchange rates can be actually confirmed in our sample. This set up can be thought as a natural panel data extension of the final Granger-causation application in Cashin, Céspedes, and Sahay (2004). In both specifications mentioned above, standard errors and p-values have been bootstrapped to take into account cross sectional correlation bias, short sample bias, and the self-induced endogeneity bias that would make inference difficult given the potential non-reliability of the estimated standard errors. To give an overview of Granger causation in a panel framework, Dumitrescu and Hurlin (2012) propose the following model:

$$y_{i,t} = \alpha_i + \Sigma_{k=1}^{K} \vartheta_i^{(k)} y_{i,t-k} + \Sigma_{k=1}^{K} \beta_i^{(k)} x_{i,t-k} + \varepsilon_{i,t} \tag{3.25}$$

with an, intercept α_i and a similar lag order k for both the objective variable and its covariates, and allowing for heterogeneous time weights $\vartheta_i^{(k)}$ and regression coefficients $\beta_i^{(k)}$. The test thus implicitly assumes there could be causality for some countries but not necessarely for all of them. As such, for a standard null hypothesis defined as $H_0 = \beta_{i1} = ... = \beta_{iK} = 0$ for all $i = 1, ..., N$, the alternative(s) will instead be $H_1' = \beta_{i1} = ... = \beta_{iK} = 0$ for $i = 1, ..., N_1$ and $\beta_{i1} \neq 0$ or ... or $\beta_{iK} \neq 0$ when $i = N_1 + 1, ..., N$. The authors thus propose to run the model in (3.25) for all cross sections, perform F-tests for each of the K linear null hyphoteses, get an individual Wald statistic, and finally compute the average statistic W, such that:

$$W = \frac{1}{N} \Sigma_{i=1}^{N} W_i \tag{3.26}$$

When the Wald statistics are i.i.d., a normally distributed Z test can also be employed:

$$\overline{Z} = \sqrt{\frac{N}{2K}}(W - K) \ \ \overrightarrow{T, N \longrightarrow \infty} \ \ N(0,1) \tag{3.27}$$

But the above definition depends on both T and N, specifically on T reaching infinity at a faster speed of convergence than N. The authors finally show how in Fixed time samples $\overline{Z}$ can be finally aproximated with:

$$\widetilde{\overline{Z}} \ \ \overrightarrow{N \longrightarrow \infty} \ \ N(0,1) \tag{3.28}$$

The testing procedure we employ is thus finally based on $\widetilde{\overline{Z}}$ from Equation (3.28).

The non Granger causality tests are reported in Table 3.13. As already explained, p-values were block-bootstrapped to ensure unbiasedness of the results due to weak cross sectional spillovers. The tests shows how only in non-energy exporting commodity countries (group IX) and in the expanded advanced country group (VI) the commodity price index would help describe the evolution of the exchange rate in the short term, while in the lower section of the table we are able to exclude "reverse causation", that is, we cannot reject the hypothesis that short term fluctuations of the exchange rate do not help us describing commodity prices in any of the groups we considered. Surprisingly, such evidence would not hold for the Cashin, Céspedes, and Sahay (2004) countries selection group (which, by all means, is restricted to fewer countries when compared to the seminal article, mainly due to the suprising lack of stored data in the IMF IFS database) and, perhaps would also not hold in the energy commodities exporters group.

Table 3.14 shows results for our rendition of the weak exogeneity test as seen in Cashin, Céspedes, and Sahay (2004), based on the error correction specification of the bilateral relationship between exchange rates and price indexes. As the distribution of the calculated coefficients is unknown to us, especially in a time series environment, we resort to block-bootrapping the standard error to get an estimate of the coefficient of error correction and of the joint significance of the lagged first differenced variables. The test basically amounts to checking for contemporaneous statistical significance of both the speed of adjustment and the time lags of the regressors. Results from this experiment show how the only group of countries which would contemporaneously present univocal

Table 3.13: Granger causality testing, 100 repetitions

H0: drcomp does not Granger-cause dlnreer_ alt					
	(I)	(V)	(VI)	$(VIII)$	(IX)
W	1.409	1.229	1.519	1.053	1.690
$\overline{Z}$	1.387	0.067	2.514**	0.151	3.550**
$p-value$	(0.180)	(0.500)	(0.040)	(0.940)	(0.040)
$\widetilde{\overline{Z}}$	1.329	0.627	2.421**	0.120	3.436**
$p-value$	(0.180)	(0.530)	(0.050)	(0.950)	(0.040)

H0: dlnreer_ alt does not Granger-cause drcomp					
	(I)	(V)	(VI)	$(VIII)$	(IX)
W	1.359	1.577	1.174	1.453	0.993
$\overline{Z}$	1.216	1.681	0.842	1.280	-0.035
$p-value$	(0.260)	(0.110)	(0.510)	(0.240)	(0.990)
$\widetilde{\overline{Z}}$	1.161	1.622	0.778	1.229	-0.086
$p-value$	(0.290)	(0.110)	(0.550)	(0.270)	(0.970)

Granger causation tests for group I, V, VI, VIII and IX. Lag selection carried out through AIC, from a maximum of three lags per cross section. Block-bootstrap was carried out with ten thousands repetitions.

Table 3.14: Weak causality testing, 10000 repetitions

H0: lnreer_alt does not adjust to the equilibrium with rcomp					
	(I)	(V)	(VI)	$(VIII)$	(IX)
ECT	-0.022**	-0.035**	-0.026**	-0.012	-0.021**
$B.s.e.$	(0.004)	(0.007)	(0.008)	(0.015)	(0.003)
χ^2	7.840*	4.080	3.470	3.100	7.270*
$p-value$	(0.050)	(0.253)	(0.325)	(0.376)	(0.064)

H0: rcomp does not adjust to the equilibrium with lnreer_alt					
	(I)	(V)	(VI)	$(VIII)$	(IX)
ECT	0.006	-0.002	-0.010	-0.012	-0.010*
$B.s.e.$	(0.007)	(0.014)	(0.008)	(0.015)	(0.005)
χ^2	20.070*	2.680	3.900	3.150	4.530
$p-value$	(0.001)	(0.444)	(0.2722)	(0.368)	(0.210)

Bootstrapped standard errors in parentheses. Lag selection carried out through AIC.

short run Granger causation, going from commodity prices to the exchange rate, and univocal long run adjustment to the equilibrium, is the group of non energy exporting commodity countries and perhaps (some) of the countries originally selected by Cashin, Céspedes, and Sahay (2004). Although we have proven so, this result is far from being a decisive confirmation of causality: on the contrary, some feedback can be seen in the long run adjustment of the non energy exporting commodity countries and some short run fluctuation feedback can also be seen in the Cashin, Céspedes, and Sahay (2004) countries we had available for the analysis. Such evidence suggests that weak causation going from commodity prices to the exchange rate is still present in emerging commodity exporting countries, although short term predictability based on the panel Granger-causation test could not be confirmed. In substance, no group would clearly present contemporaneously a significant (and meaningful) short run causal adjustment together with short term Granger-cause predictability. This finally brings forth the exigence to rivisit the discretional export threshold which have for a long period being considered as the determinant of the relationship between commodity prices and exchange rates.

3.8 Conclusions

This chapter aims to study the impact of leading commodity prices long run term trends in a panel of developed and developing countries. Through the use of new import weights and modern averaging panel techniques, which take both cross section heterogeneity and possible cross section error dependence into account, our analysis shows that introducing a commodity price index in a cointegrating relationship with the real effective exchange variable radically changes the long run impact of commodity prices on the exchange rate. Our estimates show that not only elasticities have dramatically decreased when compared to past literature, but that Granger causation of the exchange rate given by commodity price fluctuations in exporting countries does not substantially depend anymore on a relative export commodity weight of 50%, and thus cannot be generalized for the emerging exporting country group from Cashin, Céspedes, and Sahay (2004).

Chapter 4

A threshold analysis on commodity driven exchange rate convergence and adjustment

4.1 Introduction

This chapter aims to contribute to the existing literature on the real exchange rate convergence by extending current and past literature related to nonlinear behavior of the exchange rate, investigating sources of short run deviation patterns through a threshold transition panel estimation with a commodity price index as a threshold variable, thus bringing past theory models one step forward in terms of representing the Heckscher commodity point theory. We contribute to the exchange rate equilibrium literature by addressing the matter in the light of a short term model motivated by arbitrage adjustment conditions and underlining the validity and the existence of commodity points. This last point, related to the choice of different selection methods, will probably be fundamental in the development of further studies. We use a new commodity price index devised using the most updated trade weights at our disposal for a group of more than forty commodities across a panel of seventy countries. The objective is extending the contents of the previous chapter and following up the methodology in chapter three by evaluating whether exchange rate fluctuations would depend on regime nonlinearity other than being influenced by cross section non-linearity (that is, parameter heterogeneity and eventual cross section spillovers, which we analyzed in the last chapter).

This chapter is organized as follows: Sections 4.1 and 4.2 underline the contributions and motivations of the chapter. Subsection 4.2.1 offers an empirical literature review while Subsection 4.2.2 introduces and contextualizes the threshold regression models we employ in the analysis. Section 4.3 explains the database sources and variables; unit root and cointegration tests are shown in Section 4.4; Section 4.5 presents the linearity and homogeneity tests; results for the whole panel are available in section 4.6. As for group results, they are explained in detail in Section 4.7: Subsection 4.7.1 digs into the details of the regime switching models, Subsection 4.7.2 shows the linearity and homogeneity test results, while finally Subsection 4.7.3 illustrates the results of the threshold model regressions applied to each group. Section 4.8 contains a robustness check obtained by substituting the exchange rate definition with an alternative one. Section 4.9 concludes the chapter.

4.2 Motivations

4.2.1 Empirical literature review and motivations

In the 2000s, evidence of a long run relationship between real exchange rates and commodity prices had found consistent, albeit country study limited, confirmation in the seminal works of Chen and Rogoff (2003b), who first tested such relationship for a group of five advanced countries, and Cashin, Céspedes, and Sahay (2004), who switched the attention to commodity producing countries. In an attempt to look for the existence of possible long run relationships, modern literature presents a relevant scarcity of analyses proving how the exchange rate would react and readjust to disequilibria in the short term according to commodity price behavior. The most recent evidence of the nexus between commodity prices and real exchange rates can be found in Coudert, Couharde, and Mignon (2015), where the authors made an attempt at modelling an asymmetric model in a panel data framework which resembled time series smooth-transition models but considered as a threshold variable the residuals of a regression between a real exchange rate measure and a set of commodity terms of trade.

On a parallel plane of literature, seminal works such as Obstfeld and Taylor (1997b) and Michael, Nobay, and Peel (1997) focused on nonlinearity as a matter of regime switching other than parameter heterogeneity. Both the latter and the former authors would use discrete and continuous autoregressive modelling to identify a structure where parameter heterogeneity gravitates around an attractor or an attraction band, revamping the commodity point theory and the ideal existence of imperfect arbitrage. In a later stage, Baum, Barkoulas, and Caglayan (2001b) and Imbs, Mumtaz, and Ravn (2003) respectively studied Purchasing Power Parity deviations within an ESTAR (exponential smoothing threshold autoregression) framework in seventeen advanced countries, finding evidence of significant deviations from the parity at the national level, and exchange rate nonlinear behavior at the sectorial level, underlying the existence of conditional mean reversion in the exchange rate path in some economic sectors.

With Béreau, Villavicencio, and Mignon (2010), Coudert, Couharde, and Mignon (2015) and Allen, McAleer, Peiris, and Singh (2016a), the two strands of literature (namely the nonlinear autoregressive and the commodity dependence related one) finally found some common soil: thanks to modern developments in panel data techniques, the former authors attempted at modelling oil prices as a form of exogenous threshold variable in a behavioral specification of a panel of multilateral real exchange rates, while the latter made use of all the currently available methods of nonlinear time series estimation[1] to revisit the topic of exchange rate nonlinear behavior.

Our contribution to current literature involves following González, Teräsvirta, Van Dijk, and Yang (2017) to estimate a smooth transition model for various group of countries and verifying to which extent nonlinearity in the exchange rate behavior can be identified by a commodity driven regime switching behavior (arbitrage) other than depend on absolute or relative deviations from a behavioral equilibrium or be part of the natural slope heterogeneity which is proper of each cross section employed in the analysis.[2] In a sense, an answer to this issue of representation would be an answer to the exchange rate low speed of return puzzle: is a multiple regime model exchange rate return to equilibrium faster than in an heterogeneous slope set up?

[1] Including, in a similar fashion to chapter two, exponential, logistic models and a simple neural network specification.

[2] As we have seen in chapter three with the heterogeneous panel estimates we have calculated with the CCEMG (common correlated effect mean group estimator) and AMG (augmented mean group estimator).

4.2.2 On the impact of volatility on the exchange rate: the smooth threshold regression

Estimation of real exchange rate models has been carried out across literature through different choices for controls, mainly related to theoretical models such as the flexible and monetary ones, or to behavioral definitions like in Macdonald (1998) and Macdonald (2000). Specifically, much attention has been put in the past in deriving a terms of trade variable and consequently a variety of price indexes, which would enter the exchange rate equation as fundamental controls. Bodart, Candelon, and Carpantier (2015b) and Bodart, Candelon, and Carpantier (2012) suggest an interesting change in such set up allowing for the price of a selection of leading exported commodities to enter the exchange rate equilibrium equation. This chapter tries to capture the relationship between commodity prices and real exchange rates by updating the commodity price index variable created by Cashin, Céspedes, and Sahay (2004) with a set of newly updated price weights, which we sourced from the same authors.[3] We thus aim at uncovering the relationship between the variation in the real exchange rates, and the variation in its fundamentals, with a specific focus on the price of exported commodities. Our analysis entails a smooth transition model adapted for longitudinal data as in González, Teräsvirta, Van Dijk, and Yang (2017), based on the discrete regime switching one by Hansen (1999b). The generic form of the model fits a logistic function with one or two thresholds and as such describes two external converging regimes and possibly (but not necessarily) a non converging, unit root behaving central transitional one:

$$\Delta y_{i,t} = \alpha_i + \lambda_t + \beta_2^{'} \Delta X_{i,t} + (\beta_2^{''} \Delta X_{i,t}) * G(s_{i,t}; \gamma; c) + \varepsilon_{i,t} \tag{4.1}$$

where $y_{i,t}$ is once again the targeted real exchange rate, $X_{i,t}$ represents our set of variables co-moving with the real exchange rate, which includes our newly built export commodity price weighted index, and finally the transition variable $s_{i,t}$, with slope γ and cut-off value equal to c. Following González, Teräsvirta, Van Dijk, and Yang (2017),

[3] For an overview of the theoretical relationship between exchange rates and Commodity Prices, see chapters two and three of this thesis.

the logistic function will be equal to:

$$G(s_{i,t};\gamma;c) = \left\{1 + e^{\left[-\gamma\Pi_{j=1}^{m}(s_{i,t}-c_j)\right]}\right\}^{-1} \tag{4.2}$$

where γ represents the slope of the transition and c the value of the threshold.

In particular, with the self exciting smooth threshold autoregression model (SESTAR) we intend to target the first difference of the real effective exchange rate ($\Delta REER$) as a threshold variable to account for unconditional, albeit regime driven, return to the mean of the exchange rate process regardless of the absolute size of misalignment, whether the exchange rate we analyze is over or undervalued. As variables in the model are assumed to be stationary, we will work in first differences when needed. Notice that being the model a self exciting one, the value of the transition function G will be governed by a threshold variable $s_{i,t}$ basically equivalent to the objective variable of the function, so that $\Delta y_{i,t} = s_{i,t}$. A compact notation for the self exciting mode is shown below, in Equation (4.3).

$$\Delta y_{i,t} = \alpha_i + \beta_i \Sigma_{i=1}^{k} \Delta y_{i,t-1} * G(s_{i,t};\gamma;c) + \varepsilon_{i,t} \tag{4.3}$$

Together with the SESTAR application, which matches closely previous literature such as Michael, Nobay, and Peel (1997), we choose to test for additional specifications in order to capture the mechanisms underlying the deviations. As such, we introduce weakly exogenous covariates, which we have selected taking into account the equilibrium definition of the real exchange rate common in the literature, and use them to fit multiple regimes in a simple error correction model for the exchange rate, devising a panel cointegrating smooth transition regression model (PCSTR), which closely follows the Béreau, Villavicencio, and Mignon (2010) and Coudert, Couharde, and Mignon (2015) specifications:

$$\begin{aligned} \Delta y_{i,t} \quad = \quad & \alpha_i + \beta_0^{'} \Delta y_{i,t-1} + \beta_1^{'} ECT_{i,t-1} + \beta_2^{'} \Delta X_{i,t} + & (4.4) \\ & [(\beta_0^{''} \Delta y_{i,t-1} + \beta_1^{''} ECT_{i,t-1} + \beta_2^{''} \Delta X_{i,t}) * G(s_{i,t};\gamma;c)] + \varepsilon_{i,t} & (4.5) \end{aligned}$$

where every variable will assumedly be $I(0)$. Such model, if cointegration between $y_{i,t}$

and $X_{i,t}$ is nowhere to be found, would imply $\beta_1^{'} = 0$ and collapse to the standard panel smooth transition regression model (PSTR) with exogenous covariates:

$$\Delta y_{i,t} = \alpha_i + \zeta^{'} \Delta X_{i,t} + [(\zeta^{''} \Delta X_{i,t}) * G(s_{i,t}; \gamma; c)] + \varepsilon_{i,t} \tag{4.6}$$

Finally, if the slope of the transition function tends to infinity, $\gamma \longrightarrow \infty$, the regime jumps are discrete and no economic operator is inactive, and we are back to Hansen (1999b) panel threshold regression (PTR):

$$\Delta y_{i,t} = \alpha_i + \zeta^{'} \Delta X_{i,t} + [(\zeta^{''} \Delta X_{i,t}) * I(s_{i,t})] + \varepsilon_{i,t} \tag{4.7}$$

$$I = \begin{cases} 0 \; for \; s_{i,t} < c \\ 1 \; for \; s_{i,t} \geq c \end{cases} \tag{4.8}$$

It is worth pointing out that Equation (4.8) establishes a single value per threshold. However, although the amount of regimes we have tested for is equal to two, the number of threshold values we allowed for during the tests were either one or two, with $c_1 < c_2$, such that, under unit root or quasi-unit root behavior between c_1 and c_2, we would be entitled to refer to the model as a band threshold regression model, with a literal "inaction band" inside it.

Following the laws of arbitrage, the assumption made by Heckscher on the existence of commodity points, and the fact that economic agents are supposed to take some time to adjust to economic news, we will set $s_{i,t}$ to be equal to the first difference of our commodity price index, $\Delta COMM$. This would suggest the existence of convenient arbitrage opportunities and the existence of a unit root like interval where the benefit of real or nominal arbitrage are non existent for the economic operators. As a mean of uncovering how much productivity differentials matter in the speed of adjustment, we will then choose as a threshold the logaritmic first difference of per capita real GDP, ΔGDP. We will also employ the error correction term itself, ECT, as a mean of exploring whether or behavioral definitions still apply to the exchange rates and are somehow more useful than the self adjusting mechanism entailed by the use of $\Delta REER$. With that, and taking into account differences between groups in a panel of emerging and advanced countries, we will be able to infer on what might actually be the driving forces of arbitrage, whether a sensible difference exists between country groups, and how such way of capturing nonlinearity in the exchange rate differs with respect to allowing for long run parameter heterogeneity and spatial spillovers as seen in the previous chapter.

4.3 Database and variables

We take advantage of the Cashin, Céspedes, and Sahay (2004) study to build from scratch an updated commodity price index using updated trade weights, which we have already made use of in the previous chapter. Our commodity (export) price Index is as such:

$$NCOMP_{i,t} = \exp\left[\sum_{k=1}^{K} Weight_{i,k} * \ln(price_k)\right]_t \tag{4.9}$$

and:

$$COMM_{i,t} = \frac{\exp\left[\sum_{k=1}^{K} Weight_{i,k} * \ln(price_k)\right]_t}{MUV_t} \tag{4.10}$$

where $NCOMP_{i,t}$ represents the nominal commodity price index at time t within country i, $COMM_{i,t}$ represents the real commodity price index at time t within country i, $Weight_k$ is the average 1999-2004 average export weight of commodity k on the total of exports of goods of country i, $price_k$ is the international price of exported commodity k, and, finally, MUV_t represents the Manufacture Unit value index at time t used to deflate the Nominal prices $NCOMP_{i,t}$. Obviously, $\sum_{k=1}^{K} Weight_{i,k} = 1$.

As for the real effective exchange rates, we employ two different definitions of CPI based real effective exchange rate from the IFS (international financial statistics) of the international monetary fund (IMF) and, as a main benchmark, the recent alternative real effective exchange rate ($REER$) by Darvas (2012). The $REER$ definition by Darvas (2012), would be:

$$REER_{i,t} = \frac{NEER_{i,t} * CPI_{i,t}}{CPI_t^F} \tag{4.11}$$

where $REER_{i,t}$ is the real exchange rate of country i at time t evaluated against a basket of foreign currencies, $CPI_{i,t}$ is the consumer price index of country i at time t, and $NEER_{i,t}$ is the nominal exchange rate of country i in period t and finally CPI_t^F represents a weighted average of all the CPI's from trade partners for every period t.

Both the price index and the exchange rates are the geometrical averages of the country i's trade partners' CPI indexes and country i's bilateral exchange rates with its trade partners:

$$NEER_{i,t} = \prod_{f=1}^{N} NER_{f,t}^{(w_i)} \tag{4.12}$$

$$CPI_t^F = \prod_{f=1}^{N} CPI_{f,t}^{(w_i)} \tag{4.13}$$

where $NER_{f,t}^{(w_i)}$ is the bilateral exchange rate of country i with respect to foreign partner f, weighted by the country weight w_i, and $CPI_{f,t}^{(w_i)}$ represents the consumer price index of trade partner f at time t weighted by home country weight w_i. Once again, $\sum_{i=1}^{N} w_i = 1$, and the weights are time invariant, consistently with the previous chapter.

As we have already stated, our measure for the real exchange rate comes from one of the newest sources available, namely Darvas (2012), which allowed us to build a panel of seventy-two countries. The variables we have taken from the World bank's world development indicators database (WB WDI) would be instead net foreign asset position (NFA), real per Capita GDP (GDP) and, finally, the share of government spending over GDP (GOV). Following data availability and to keep our results comparable to previous literature, our database ranges monthly from 1995 to 2012.

4.4 Unit root test and cointegration

In order to fit a smooth transition regression model, we had to make sure that all the variables of the panel were nonstationary. That would allow us to work comfortably with their first difference and, if cointegration were present, an error correction term. All the variables ended up being integrated of order one. As we checked for levels and first differences to check for stationarity in variance, we followed Dickey and Pantula (2002). Results of the two test we employed, Maddala and Wu (1999) first generation unit root test and the Pesaran (2007) CIPS test, a second generation test which would account for possible cross-section correlation, are reported in this section with up to six lags and different deterministic structures. The results of the tests are identical to the ones we obtained in the previous chapter, so that we invite the reader to refer to that

Table 4.1: Unit root tests

$Variable$	$Maddala\ and\ Wu$ (1999)	$P - value$	$Pesaran$ (2007)	$P - value$
$REER$	156.187	0.166	-0.097	0.461
$COMM$	56.673	1.000	0.568	0.715
GDP	129.817	0.720	8.091	1.000
NFA	132.088	0.671	-1.387	0.083
GOV	139.166	0.504	0.420	0.663

Unit root tests in levels, null hypothesis: the variable has a unit root. Up to six lags tested. The alternative for the Pesaran test includes a trend.

part of the thesis for a more detailed overview of the analysis. Results are visible in Table 4.1, which we sourced from chapter three.

In Table 4.2 we also report a set of necessary cointegration tests, also sourced from chapter three. We employ the Westerlund (2007) cointegration tests borrowing the methodology from Blackburne and Frank (2007), which allows to test for the statistical significance of an error correction specification with both homogeneous and non homogeneous alternative hypotheses. Probability values were obtained through residual bootstrapping and sourced by the underlying empirical distribution to take into account the impact of possible cross-section correlation and the analysis is carried out with a top-down strategy. Overall cointegration of the panel, indicated by the P_t and P_a tests in Table 4.2, is confirmed, exception done for the result in the G_a group test for the commodities. Such final result allowed us to set up and estimate a linear error correction specification and use the resulting residuals as an error correction term for our threshold specifications.

Table 4.2: Cointegration tests

$REER$	G_t	$P-value$	G_a	$P-value$
$COMM, GDP, NFA, GOV$	-2.874	0.020	-15.655	0.030
$COMM, GDP, NFA$	-2.613	0.010	-13.034	0.010
$COMM, GDP$	-2.322	0.040	-10.661	0.040
$COMM$	-2.119	0.040	-8.567	0.070
$REER$	P_t	$Pvalue$	P_a	$Pvalue$
$COMM, GDP, NFA, GOV$	-23.698	0.040	-13.928	0.010
$COMM, GDP, NFA$	23.094	0.000	-12.706	0.000
$COMM, GDP$	-20.995	0.010	-10.507	0.020
$COMM$	-18.486	0.010	-8.291	0.010

Cointegration tests. Robust p-values are bootsptrapped one hundred times. Constant included. The panel tests both have a homogeneous null of non stationarity. The group tests have a heterogeneous null of non-stationarity in variance: at least one section show cointegration.

4.5 Nonlinear tests and short run volatility

We now focus on the short run adjustment to investigate nonlinearities of the exchange rate in a panel framework. We resort to two differently distributed Lagrangean Multiplier tests as in González, Teräsvirta, Van Dijk, and Yang (2017) to check for the existence of possible regime switches and test for the possibility of detecting convergence around an inaction threshold/band. Such tests would basically compare a sequential null hypothesis, going from linearity to up to two thresholds, with all the other possible alternatives, allowing us to infer on threshold nonlinearity of the exchange rate behavior with respect to the chosen threshold variables. We mention that, following these authors, the test would require substituting the logistic function $G(s_{i,t};\gamma;c)$ seen in Equation (4.1) with its first order Taylor expansion centered around a slope value of 0 ($\gamma = 0$). After reparametrization, Equation (4.1) would then be equivalent to:

$$y_{it} = \mu_i + \beta_0^{'*} x_{it} + \beta_1^{'*} x_{it} q_{it} + ... + \beta_m^{'*} x_{it} q_{it}^m + u_{it}^* \quad (4.14)$$

In Equation (4.14), testing for $H_0 : \beta_1^{'*} = ... = \beta_m^{'*} = 0$ is thus equivalent to the null hypothesis of the original model, $H_0 : \gamma = 0$, where again m represents the number of regimes to test for and γ the slope of the transition. The results we will see in this section offer two kinds of renditions of the test, the first with a χ^2 distributed version of the statistic and the second with the statistic distributed like an F, which is assumed to be more reliable in finite samples.

As we have stated previously, in the complete panel section, we choose as threshold variables the first difference of the commodity index we built, the variation in real GDP per capita and the real exchange rate first difference itself, in the spirit of Obstfeld and Taylor (1997b) and Taylor, Peel, and Sarno (2001a). Results are visible in Table 4.3.

Results from the tests allow us to determine a regime switch with a single threshold value for both the commodity prices and the exchange rates and a two regime, two threshold value (previously defined c_1 and c_2) switch for the GDP.

Table 4.3: Linearity and regime number tests

$\Delta COMP$ lin *vs* m	$LM\chi^2$	$P-value$	LM_F	$P-value$
0 *vs* 1	33.300	0.001	5.515	0.000
0 *vs* 2	54.200	0.000	4.487	0.000
1 *vs* 0	33.300	0.001	5.515	0.000
2 *vs* 1	20.960	0.002	3.471	0.002

$\Delta REER$ lin *vs* m	$LM\chi^2$	$P-value$	LM_F	$P-value$
0 *vs* 1	2966.000	0.000	491.200	0.000
0 *vs* 2	6059.000	0.000	501.600	0.000
1 *vs* 0	2966.000	0.000	491.200	0.000
2 *vs* 1	4002.000	0.000	662.600	0.000

ΔGDP lin *vs* m	$LM\chi^2$	$P-value$	LM_F	$P-value$
0 *vs* 1	25.330	0.001	4.195	0.001
0 *vs* 2	46.050	0.000	3.812	0.000
1 *vs* 0	25.330	0.001	4.195	0.001
2 *vs* 1	20.760	0.002	3.437	0.002

Nonlinearity LM and F tests comparing sequential null alternatives up to two possible regimes.

4.6 Whole panel results

Results for the selected specifications with the single threshold commodity index, the single threshold real GDP per capita, and the double threshold real exchange rate models are placed in Tables 4.4, 4.6 and 4.5 respectively. Our linear benchmark single equation error correction model, from which we derived our error term estimates for the nonlinear analysis, closely follows Coudert, Couharde, and Mignon (2015) application and after lag selection amounts to:

$$\Delta REER_{i,t} = \alpha_i + \phi_1 \Delta ECT_{i,t-1} + \beta_1 \Delta REER_{i,t-1} + \beta_1 \Delta COMM_{i,t} + \quad (4.15)$$

$$+\beta_1 \Delta GOV_{i,t} + \beta_1 \Delta GDP_{i,t} + \beta_1 \Delta NFA_{i,t} + \varepsilon_{i,t} \quad (4.16)$$

Results from the estimate show statistically significant threshold values for both the logistic models with one threshold and the one with two thresholds. As implied by the value of the estimated slope of the function, such transition presents varying levels of subtlety, ranging from 27300, an almost instantaneous transition, to 1134, a much more sloped one. Our aggregate results present a very steep transition, visible in Figures 4-1 and 4-2. Across all possible specifications, we do see that the two most extreme regimes presents a statistically significant and converging error correction value, while the nonlinear central regime stays nonsignificant across all three specifications. The values of the speed of error correction range between -0.0414 to -0.0157, and on average are consistently higher than their linear equivalents, as estimated in the previous chapter. As we go from one most extreme regime to the other, we do notice how the adjustment tends to behave, in the overall panel, very symmetrically, as probably would be expected of a world aggregate merging a heterogeneous spectrum of countries.[4]

On average, productivity differentials appear to be the most likely important determinants of short run adjustment, with an average adjustment coefficient of 0.02855 once the per capital real GDP variation is used as a threshold variable. Furthermore, the short run relationship appears to be dominated by adjustment to the long run equilibrium and by its past values, as inferred by the always consistently positive and

[4] In the following paragraph we will illustrate group estimates of the model. Following Béreau, Villavicencio, and Mignon (2010), we would expect to see a more asymmetric behavior.

Table 4.4: PLSTR, threshold variable COMM

	Regime I	*Non-linear*	*Regime II*
ECT	-0.0195	-0.0100	-0.0295
	(0.0031)	(0.0087)	(0.0067)
$\Delta REER_{t-1}$	0.1918	0.8996	0.2817
	(0.0453)	(0.0849)	(0.0523)
$\Delta COMM$	0.0025	-0.0031	-0.0006
	(0.0064)	(0.0071)	(0.0040)
ΔGOV	0.0053	0.0004	0.0009
	(0.0002)	(0.0003)	(0.0003)
ΔGDP	0.0209	0.0009	0.2171
	(0.0132)	(0.0162)	(0.0175)
ΔNFA	-0.0006	-0.0040	-0.0046
	(0.0021)	(0.0025)	(0.0016)
$\widehat{c}_1$	0.0153		
	(0.0012)		
$\widehat{c}_2$			
$\widehat{\gamma}$	1134		
	(559.3)		

Standard errors in parentheses. Lag selection carried out through AIC. the threshold vaule is indicated by the letter c, while γ represents an estimate of the slope of the transition.

significant coefficients of its lagged values. Such evidence does suggest the importance of both arbitrage opportunities and absolute size of over/undervaluation in exchange rate determination, but cannot yet be considered valid to infer on short term causation or to understand differences among countries whose export structure differ in relation to the weight of their commodity sector. We will see what that entails in the next section.

Table 4.5: PLSTR, threshold variable REER

	Regime I	*Non-linear*	*Regime II*
ECT	-0.0215	-0.0030	-0.0245
	(0.0128)	(0.0226)	(0.0104)
$\Delta REER_{t-1}$	0.2004	0.0118	0.2122
	(0.0507)	(0.0790)	(0.0344)
$\Delta COMM$	0.0011	-0.0022	-0.0012
	(0.0050)	(0.0040)	(0.0043)
ΔGOV	0.0006	0.0001	0.0007
	(0.0005)	(0.0009)	(0.0006)
ΔGDP	-0.0886	0.2557	0.1671
	(0.0264)	(0.0385)	(0.0204)
ΔNFA	0.0000	-0.0057	-0.0057
	(0.0032)	(0.0069)	(0.0042)
$\widehat{c_1}$	0.0032		
	(0.0002)		
$\widehat{c2}$			
$\widehat{\gamma}$	13220		
	(2349)		

Standard errors in parentheses. Lag selection carried out through AIC. The threshold value is indicated by the letter c, while γ represents an estimate of the slope of the transition.

Table 4.6: PLSTR, threshold variable GDP

	Regime I	*Non-linear*	*Regime II*
ECM	-0.0157	-0.0257	-0.0414
	(0.0045)	(0.01857)	(0.0148)
$\Delta REER_{t-1}$	0.2978	-0.1456	0.1522
	(0.0180)	(0.0533)	(0.0453)
$\Delta COMM$	0.0026	-0.0027	0.0000
	(0.0048)	(0.0063)	(0.0054)
$DGOV$	0.0059	0.0003	0.0009
	(0.0003)	(0.0006)	(0.0005)
ΔGDP	0.0285	-0.0165	0.0120
	(0.1496)	(0.0220)	(0.0196)
ΔNFA	-0.0005	-0.0039	-0.0044
	(0.0022)	(0.0033)	(0.0020)
$\widehat{c}_1$	0.0542		
	(0.0005)		
$\widehat{c}_2$	-0.0208		
	(0.0040)		
$\widehat{\gamma}$	27300		
	(7025)		

Standard errors in parentheses. Lag selection carried out through AIC. The threshold value is indicated by the letter c, while γ represents an estimate of the slope of the transition.

Table 4.7: Selected groups

Group	*Authors*	*Countries*	*Type*
I	*Cashin et al. (2004)*	24	commodity countries
II	*Ricci et al. (2013)*	21	adv. countries, IMF
III	*Ricci et al. (2013)*	10	adv. countries, non EU, IMF
IV	*Ricci et al. (2013)*	11	adv. countries, EU, IMF
V	*Ricci et al. (2013)*	17	nearly adv., IMF
VI	*(1-III)+IMF new*	47	non advanced countries
VII	*III+IMF new*	23	adv. countries, IMF
VIII	*Bodart et al. (2013)*	16	energy major exp., IMF
IX	*Bodart et al. (2013)*	54	non energy major exp., IMF

Group estimates. The WCED IMF we employed to build groups VIII and IX defines major exporters those having a comodity export weight equal or more than 20 per cent of the overall export volume.

4.7 Group results

4.7.1 Exploring the regime-switching hypothesis

We now look at the overall results for a sequence of national groups divided in advanced, emerging, commodity dependent and energy and non energy exporting countries. The groups have been selected based on their availability and their presence in seminal literature studies. In particular, we have nine main groups for which we shall test the nonlinearity hypothesis.

In Table 4.7 we present the origin and definition of the groups we tested for in the group result section. In particular, *group I* contains the following countries: Australia, Bolivia, Brazil, Central African Republic, Canada, Chile, Cote d'Ivoire, Cameroon, Colombia, Costa Rica, Ghana, Morocco, Mexico, Malaysia, Nicaragua, Norway, New Zealand, Pakistan, Philippines, Paraguay, Tunisia, Uganda, Uruguay and South Africa. *Group II* contains Australia, Austria, Belgium, Canada, Switzerland, Germany, Denmark, Spain, Finland, France, United Kingdom, Greece, Ireland, Italy, Japan, Netherlands, Norway, New Zealand, Portugal, Sweden and United States. *Group III* contains Australia, Canada, Switzerland, Denmark, United Kingdom, Japan, Norway, New Zealand, Sweden and United States. *Group IV* contains Austria, Belgium, Germany, Spain, Finland, France, Greece, Ireland, Italy, Netherlands and Portugal. *Group V* contains Brazil, Chile, China, Colombia, Czech Republic, Hungary, Morocco, Mexico, Malaysia, Pakistan, Philippines, Poland, Russia, Singapore, Slovak Republic, Venezuela and South Africa. *Group VI* contains: Bulgaria, Bahrain, Belize, Bolivia, Brazil, Central African republic, Chile, China, Cote d'Ivoire, Cameroon, Colombia, Costa Rica, Cyprus, Czech Republic, Dominican Republic, Algeria, Gabon, Georgia, Ghana, Gambia, Guyana, Hungary, Iran, Latvia, Morocco, Moldova, Mexico, Macedonia, Malta, Malaysia, Nigeria, Nicaragua, Pakistan, Philippines, Poland, Paraguay, Romania, Russia, Saudi Arabia, Singapore, Slovak Republic, Tunisia, Uganda, Ukraine, Uruguay, Venezuela and South Africa. *Group VII* contains: Australia, Austria, Belgium, Canada, Switzerland, Germany, Denmark, Spain, Finland, France, United Kingdom, Greece, Ireland, Iceland, Israel, Italy, Japan, Netherlands, Norway, New Zealand, Portugal, Sweden and United States. *Group VIII* contains: Bahrain, Bolivia, Cote d'Ivoire, Cameroon, Colombia, Algeria, Gabon, Ghana, Guyana, Iran, Mexico, Nigeria, Norway, Russia and Saudi Arabia. Finally, *Group IX* contains: Australia, Austria, Belgium, Bulgaria, Belize, Brazil, Central African Republic, Canada, Switzerland, Chile, China, Costa Rica, Cyprus, Czech Republic, Germany, Denmark, Dominican Republic, Spain, Finland, France, United Kingdom, Georgia, Gambia, Greece, Hungary, Ireland, Iceland, Israel, Italy, Japan, Latvia, Morocco, Moldova, Macedonia, Malta, Malaysia, Nicaragua, Netherlands, New Zealand, Pakistan, Philippines, Poland, Portugal, Paraguay, Romania, Singapore, Slovak Republic, Sweden, Tunisia, Uganda, Ukraine, Uruguay, United States and South Africa.

Considering the importance of possible spatial spillovers when clustering groups of cross section units, we decided to infer on the possible existence of nonlinear threshold behavior taking into account statistics which would be resistant to spheric disturbances and cross sectional correlation.

We remind that in the linearity checks we employed we are testing for a nonlinear logistic structure as an alternative to the null of linearity. How much *logistic* that alternative could be, how sloped the transition is, that is up to the estimations to tell. To account for standard spherical disturbances as well as cross sectional correlation, we coupled the standard within, HAC robust tests with the very same tests with probability values which have been wild bootstrapped following the procedure described in González, Teräsvirta, Van Dijk, and Yang (2017). In our test for smooth transitions, we have framed two possible group of thresholds: first, a group composed of the real exchange rate itself, the residuals of the level regression, and the commodity price indexes, which can be assumed to be "compounded thresholds", as they are either a linear combination of our controls or a weighted index variable; second, a group composed of each singular exogenous control, our "secondary thresholds", which might contribute not just to simple volatility but also to regime switches in the adjustment path of the exchange rate.

Overall, a total of fifty-four different specifications were tested, with results showing the presence of valid inference only in the case of very sharp transitions, that is whether the slope of the $\gamma = e^{\delta}$ factor approached infinity.

4.7.2 Linearity and homogenity test results

This section shows the results for the homogeneity tests. As we have briefly explained in the last paragraph, we have chosen to divide the tests according to two threshold groups: a "compounded" one, made from either indexes we have estimated or measures we have calculated (Table 4.8) and a "secondary" one, made of the controls we used up until now, which form part of the literature on the behavioral definition and adjustment of the exchange rate (Table 4.9). Results from Table 4.8 highlight two findings: first, variations of the exchange rate dominate the transition between different states of arbitrage. Being an unconditional transition, as the threshold variable is nothing but a lag of the objective variable, this applies to eight out of the nine country groups, exception done for oil exporters, and does allow us to generalize the common knowledge that absolute size matters in the determination of the adjustment speed of the exchange rate. In a sense, such result represents the long awaited generalization of the results achieved by Obstfeld and Taylor (1997b) in a time series setting, where regardless of the lack of a specific structural form for the Commodity Points theory (which we offer thanks to our choice of $\Delta COMM$ as a threshold) the exchange rate was found to be by itself useful in describing threshold behavior.

Another result, perhaps starkingly different from what we would have somehow expected, is that not only, in contrast with the result obtained by Béreau, Villavicencio, and Mignon (2010), the exchange rate is characterized by nonlinearities in advanced as much as in emerging countries, but such nonlinearity also appears[5] driven by the variation in commodity prices, which lead the nonlinear adjustment in the avanced country groups II, III, IV and VII.

Table 4.9 contains the contribution of the behavioral controls to threshold variation. The variation of GDP, net foreign asset position, and government spending share on GDP were all considered as suitable candidates to identify sources of variation and adjustment of the real exchange rate, being part of the behavioral definition itself. Results show evidence of how variations in GDP per capita are indeed responsible for regime switches in five out of nine country groups. Such evidence is enough to lead us to conclude that productivity differentials affecting the relative weight of the tradable and non tradable sector through changes in the working force composition are indeed part of the drivers of the change in regime, and such effect does not appear to be limited by the composition of the country groups, as both the Cashin, Céspedes, and Sahay (2004) selection of commodity exporting country, our selection of IMF related commodity exporting countries with a twenty percent export share, as well as four more advanced country groups, appeared to point at GDP variations as the underlying source of change. Remarkably, no regime switch was found in the oil exporting group.

4.7.3 Error correction and adjustment heterogeneity

In this section we finally look at how the short term fluctuations and eventual adjustment speed and inertia are conditioned by the regime switches. Is the speed of adjustment in the extreme regimes different from any linear specification? Does the adjustment differ from one extreme regime to the other? Does an inaction band really exist around an attractor value? Is the regime switch relatively slower when compared to past literature focusing on discrete jumps, or is it more gradual? The estimation aims at inferring on the behavior of the branches of the split regression, the slope of the transition, and the speed of adjustment.

Results for the group estimates are visible in Tables 4.10, 4.11 and 4.12. In a similar fashion to the complete panel estimates, we have chosen to use the variation in the real

[5] We insist in saying "appear" as the nonlinearity tests are by themselves not sufficient to establish whether or not the threshold is statistically significant. That will be up to the estimates, in particular to the estimate of the exact threshold value c, in the next section.

Table 4.8: Homogeneity tests I

		LM_F-stat		*WCB P-values*		*LM_F alt*	
Group	$s_{i,t}$	*0 vs 1*	*0 vs 2*	*0 vs 1*	*0 vs 2*	*1 vs 0*	*2 vs 1*
I	$\Delta COMM_{i,t}$	8.187	14.490	0.31	0.84	8.178	5.022
	$\Delta REER_{i,t}$	11.950	18.690	0.03**	0.00**	11.950	12.250
	$ECT_{i,t-1}$	10.320	13.560	0.00***	0.04**	10.320	5.270
II	$\Delta COMM_{i,t}$	12.560	15.210	0,03**	0.05**	12.560	8.088
	$\Delta REER_{i,t}$	13.110	18.930	0.00***	0.00***	13.110	6.859
	$ECT_{i,t-1}$	9.449	17.450	0.19	0.20	9.449	4.425
III	$\Delta COMM_{i,t}$	5.602	9.201	0.08*	0.34	5.602	5.265
	$\Delta REER_{i,t}$	8.047	10.470	0.00***	0.01**	8.047	4.724
	$ECT_{i,t-1}$	7.606	9.844	0.19	0.24	7.606	5.649
IV	$\Delta COMM_{i,t}$	9.600	14.360	0.02**	0.00***	9.600	8.599
	$\Delta REER_{i,t}$	10.460	9.695	0.00***	0.00***	10.460	8.167
	$ECT_{i,t-1}$	9.752	11.480	0.01**	0.00***	9.752	9.308
V	$\Delta COMM_t$	2.675	12.700	0.74	0.78	2.675	5.749
	$\Delta REER_t$	7.723	14.350	0.22	0.00***	7.723	10.980
	$ECT_{i,t-1}$	8.391	13.370	0.02**	0.31	8.391	10.040
VI	$\Delta COMM_{i,t}$	6.335	15.550	0.50	0.52	6.335	9.059
	$\Delta REER_{i,t}$	9.045	22.880	0.86	0.00***	9.045	11.830
	$ECT_{i,t-1}$	8.883	13.540	0.76	0.05*	8.883	9.375
VII	$\Delta COMM_t$	10.540	15.890	0.06*	0.12	10.540	8.949
	$\Delta REER_t$	13.810	19.880	0.00***	0.00***	13.810	6.941
	$ECT_{i,t-1}$	10.010	17.920	0.20	0.25	10.010	3.306
VIII	$\Delta COMM_t$	5.307	12.520	0.62	0.71	5.307	8.725
	$\Delta REER_t$	6.845	13.540	0.56	0.26	6.845	10.900
	$ECT_{i,t-1}$	5.704	14.320	0.60	0.02**	5.704	9.082
IX	$\Delta COMM_t$	4.403	9.658	0.32	0.19	4.403	8.939
	$\Delta REER_t$	13.220	26.190	0.00***	0.00***	13.220	11.100
	$ECT_{i,t-1}$	15.160	20.010	0.03**	0.62	15.160	6.945

Linearity tests for the nine country groups. Wild clustered bootstrapped values account for heteroschedasticity, possible intra-autocorrelation and cross-sectional dependence.

Table 4.9: Homogeneity tests II

		LM_F-stat		*WCB P-values*		*LM_F alt*	
Group	$s_{i,t}$	*0 vs 1*	*0 vs 2*	*0 vs 1*	*0 vs 2*	*1 vs 0*	*1 vs 0*
I	$\Delta GDP_{i,t}$	10.360	14.500	0.09*	0.00***	10.360	9.054
	$\Delta NFA_{i,t}$	6.601	13.660	0.14	0.01**	6.601	9.315
	$\Delta GOV_{i,t}$	8.430	14.760	0.51	0.32	8.430	6.790
II	$\Delta GDP_{i,t}$	7.839	17.390	0.33	0.06*	7.839	9.283
	$\Delta NFA_{i,t}$	6.249	14.980	0.65	0.52	6.249	6.509
	$\Delta GOV_{i,t}$	8.258	13.340	0.44	0.01**	8.258	9.160
III	$\Delta GDP_{i,t}$	7.103	5.021	0.59	0.09*	7.103	6.590
	$\Delta NFA_{i,t}$	6.364	9.222	0.97	0.96	6.364	8.214
	$\Delta GOV_{i,t}$	4.505	8.551	0.66	0.15	4.505	6.257
IV	$\Delta GDP_{i,t}$	9.000	11.270	0.00***	0.00***	9.000	7.929
	$\Delta NFA_{i,t}$	6.431	11.080	0.51	0.40	6.431	7.062
	$\Delta GOV_{i,t}$	10.290	11.780	0.03**	0.00***	10.290	10.180
V	$\Delta GDP_{i,t}$	5.405	13.450	0.03**	0.00”’	5.405	12.740
	$\Delta NFA_{i,t}$	6.472	14.540	0.79	0.66	6.472	14.540
	$\Delta GOV_{i,t}$	9.973	16.730	0.30	0.20	9.973	16.730
VI	$\Delta GDP_{i,t}$	7.776	11.450	0.86	0.70	7.776	4.624
	$\Delta NFA_{i,t}$	8.930	17.950	0.13	0.15	8.930	7.230
	$\Delta GOV_{i,t}$	7.658	11.550	0.15	0.25	7.658	7.953
VII	$\Delta GDP_{i,t}$	8.638	16.170	0.49	0.07*	8.638	9.702
	$\Delta NFA_{i,t}$	6.473	15.420	0.72	0.38	6.473	6.932
	$\Delta GOV_{i,t}$	8.653	13.480	0.28	0.12	8.653	9.243
VIII	$\Delta GDP_{i,t}$	4.504	13.380	0.16	0.37	4.504	7.798
	$\Delta NFA_{i,t}$	7.206	13.430	0.29	0.35	7.206	5.186
	$\Delta GOV_{i,t}$	2.572	8.489	0.49	0.36	2.572	4.223
IX	$\Delta GDP_{i,t}$	11.170	11.710	0.00***	0.00***	11.170	1.367
	$\Delta NFA_{i,t}$	6.282	12.050	0.45	0.45	6.282	6.672
	$\Delta GOV_{i,t}$	7.488	11.850	0.26	0.45	7.488	11.850

Linearity tests for the nine country groups. Wild clustered bootstrapped values account for heteroschedasticity, possible intra-autocorrelation and cross-sectional dependence.

Table 4.10: Group results, threshold: commodity price index

Group	$\beta_1^{'}$	*Non-linear*	$\beta_1^{'} + \beta_2^{''}$	γ	c
I			LINEAR		
II	-0.019*	-0.003	-0.022*	1133*	0.007*
	(0.003)	(0.004)	(0.004)	(465)	(0.001)
III	0.017*	-0.012*	-0.029*	1014*	0.008*
	(0.050)	(0.003)	(0.006)	(704)	(0.020)
IV	-0.034*	0.026*	0.007	1049	0.008*
	(0.004)	(0.009)	(0.008)	(1935)	(0.001)
V			LINEAR		
VI			LINEAR		
VII	-0.019*	-0.003	-0.022*	1093*	0.008*
	(0.003)	(0.003)	(0.004)	(460)	(0.001)
VIII			LINEAR		
IX			LINEAR		

Standard errors are in parentheses.

Table 4.11: Group results. threshold: error correction term

Group	$\beta_1^{'}$	*Non-linear*	$\beta_1^{'} + \beta_2^{''}$	γ	c
I	-0.036*	0.016	-0.020*	725	0.011*
	(0.008)	(0.011)	(0.006)	(2234)	(0.002)
II			LINEAR		
III			LINEAR		
IV	-0.027*	-0.041*	-0.068*	576*	0.014*
	(0.005)	(0.006)	(0.007)	(216)	(0.004)
V	0.015	-0.095	-0.080*	1.005*	0.00*
	(0.050)	(0.101)	(0.052)	(0.903)	(0.00)
VI			LINEAR		
VII			LINEAR		
VIII	0.023	-0.071	-0.048*	1.005*	0.00*
	(0.059)	(0.091)	(0.033)	(0.975)	(0.00)
IX	-0.022*	0.002	-0.020*	640	0.013*
	(0.005)	(0.005)	(0.006)	(711)	(0.004)

Standard errors are in parentheses.

Table 4.12: Group results, threshold: real exchange rate

Group	$\beta_1^{'}$	*Non-linear*	$\beta_1^{'} + \beta_2^{''}$	γ	c
I	-0.014*	-0.018	-0.032*	21830*	0.002*
	(0.011)	(0.021)	(0.011)	(3670)	(0.000)
II		NO NLS CONVERGENCE			
III		NO NLS CONVERGENCE			
IV		NO NLS CONVERGENCE			
V		NO NLS CONVERGENCE			
VI		NO NLS CONVERGENCE			
VII		NO NLS CONVERGENCE			
VIII		NO NLS CONVERGENCE			
IX	-0.007*	-0.050	-0.056*	21630*	0.013*
	(0.006)	(0.020)	(0.015)	(5240)	(0.000)

Standard Errors are in parentheses.

exchange rate as a threshold (Table 4.12) to take into account how adjustment varies in an appreciated state, that is, below the estimated value of the threshold, and in a depreciated state. The choice of the second threshold, which we add in this section to follow previous literature, fell on the error correction term (Table 4.11) from the linear estimates of the previous chapter. Both of these two choices above ideally refer to the self exciting smooth transition regression models which have been employed in the past, and are theoretically linked to the existence of stronger arbitrage opportunity away from the exchange rate equilibrium/parity. Our core contribution in this section is linked to the third selected threshold: in Table 4.10 we offer an overview of the results stemming from the panel smooth transition regression model when our commodity price index is used as a threshold, acting as an exogenous source of variation embodying the commodity point theory itself.

A commonality every specification shares is an almost discrete single transition, with γ values ranging from 725 to 21830. A perhaps more interesting result is given by the threshold adjustment values c: the range of values goes from 0.002, corresponding to a devaluation of around 0.2%, to 0.013, which amounts to a devaluation of 1.3%. In general, the process of convergence appears faster and statistically more consistent in the upper regimes of our specifications, when the currencies are undervalued with respect to their behavioral rate. As we have seen in the whole panel estimates, the process of adjustment tends overall to be more effective when currencies are undervalued. Relative speeds of asymmetric adjustment tend to be higher on average than linear ones, but, as far as our analysis goes, they are lower in magnitude compared to similar exercises done in the past. As a benchmark comparison, the highest convergence found by Béreau, Villavicencio, and Mignon (2010) in their comparable exercise was -0.143, about a three-fold higher than our fastest result, -0.056.

Such behavior can be generally confirmed, but reaches its highest magnitudes when the Cashin, Céspedes, and Sahay (2004) country selection (*group I*) is considered together with the IMF commodity exporting countries selection (*group IX*) as it is evident in the two coefficients for $\beta_1^{'} + \beta_2^{''}$ in Table 4.12. Notably, *group I* demonstrates the existence of an authentic inaction band, where the transitional coefficient between $\beta_1^{'}$ and $\beta_1^{'} + \beta_2^{''}$ will be non significant and as such indicate absence of adjustment. However, the transitions shown in Table 4.12 and represented by the value of slope γ are somewhat instantaneous. The modern inaction band as such, contrary to some past literature results appear to be pretty short lived. In terms of symmetry, adjustment to parity appears to occur faster at more devaluated levels, with the highest level of relative c achieved in *group IX* confirming the tendency for emerging countries and, we might add, commodity exporting ones, to have on average a more undervalued rate (perhaps as a tendency to exhaust all possible export led gains, given the export-led structure of their economies).

Switching out attention to Table 4.11, we now wonder which country groups adjust equivalently to a shock disregarding its size, or in other words, whether there exist countries where regime switching is not conditional on the magnitude of the error correction term. As it turns out, countries making part of *groups II*, *III*, *VI*, *VII* clearly do not make the cut. In such specifications, exception done for *group VI*, we have included all possible advanced countries specifications. It appear thus that "(relative) size don't matter" for mature economies, while the magnitude of the shock[6] appears to be an important co-determinant of the adjustment in commodity exporting countries, both energy and not (groups *VIII* and *IX*), where we detect a more pronounced asymmetric behavior for energy exporting country (which tend to move more in case of devaluation) and to a nonlinear but rather more homogeneous behavior in the case of non energy exporting countries.

To give further emphasis to such evidence, a similar behavior is visible in *groups I* and *V* (that is, the Cashin, Céspedes, and Sahay (2004) selection and the Nearly Industrialized and emerging economy selection by Ricci, Milesi-Ferretti, and Lee (2013)), with *group I* being slightly more homogeneous and *group V* clearly presenting higher convergence in a state of devaluation. Remarkably, *group IV* made to the regression table albeit being composed of advanced economies. In such a case however the inner regime estimated has a negative and significant coefficient, thus negating the idea of a point/band of inertia as implied by the existence of commodity points.

We finally switch our complete attention to Table 4.10. The results for such table

[6] In other words, the implied misalignment, the distance of the actual value of the exchange rate from its behavioral fundamentals.

are notable, especially since one might assume commodity dependent countries to be generally more influenced by price index behavior than more mature countries. None of the groups of commodity exporting or emerging countries, *group I*, *V* or *IX,* presented some valid results in terms of non-linearity (at least not the kind we are exploring in this chapter through panel transition regression techniques) while advanced economies, especially *groups II*, *III*, and *VII*, those countries who were part of the Ricci, Milesi-Ferretti, and Lee (2013) selection, gave us consistent nonlinear results although with a far more subtle hint of asymmetry, with regime switches accounting for 16% to 70% more of the original value of the error correction speed. This is perhaps one of the most sensible findings of this chapter.

Table 4.13: Linearity tests, IMF exchange rate

$\Delta COMM$ threshold	LM_{χ^2}	*P-value*	LM_F	*P-value*	$HAC\ LM_F$	*P-value*
0 against 1	167.500	(0.000)*	27.740	(0.000)*	2.618	(0.015)*

Linearity tests for the whole panel, IMF exchange rate as objective variable.

4.8 Robustness check

The exchange rate in his behavioral formulation depends on its fundamentals and the way they have been measured or calculated. This implies that a unique measure of the exchange rate and of its deviations simply does not exist. Past literature has looked for alternative measures of productivity and terms of trade among others as robustness checks. We remind that all our main results were obtained by using a newly devised real exchange rate series, which we choose as it gave us the opportunity to furtherly expand our analysis by reaching up to seventy cross sections and creating relevant groups as seen in Table 4.7. Thus, other than looking on a set of alternative fundamentals, our choice is "the other way around", that is to take advantage of an alternative definition of the real effective exchange rate from the IMF to check for constancy and adherence of the results obtained with Darvas (2012) measurements at the whole panel level of aggregation, given the limitations of the IMF measures. To avoid irrelevant repetitions, we will only show the estimates of the threshold model for the commodity price index as a transition variable. Results are available in Tables 4.13 and 4.14. Linearity tests would reject the presence of linearity at five per cent, while results on threshold convergence would once more present a significant non zero threshold, an instantaneous transition. The results obtained with our benchmark real exchange rate definition underline a significant convergence in outer regimes, which gets faster for a more overvalued currency, at a slightly higher rate than what we estimated two sections ago.

Table 4.14: Panel results, IMF exchange rate

	Regime I	*Non-linear*	*Regime 2*
ECT	-0.0191	-0.0483	-0.0674
	(0.0034)	(0.2933)	(0.0267)
$\Delta REER_{t-1}$	0.1743	-0.1477	0.0266
	(0.0476)	(0.0543)	(0.0125)
$\Delta COMM$	0.0028	-0.0082	-0.0054
	(0.0074)	(0.0127)	(0.0104)
ΔGOV	0.0006	0.0002	0.0008
	(0.0002)	(0.0009)	(0.0009)
ΔGDP	0.0155	-0.0647	-0.0491
	(0.0975)	(0.0631)	(0.0661)
ΔNFA	-0.0014	-0.0078	-0.0092
	(0.0013)	(0.0070)	(0.0070)
$\widehat{c}$	0.1049		
	(0,0014)		
$\widehat{\gamma}$	1153		
	(2033)		

Panel results for the threshold model with the commodity price index variation as threshold variable and the alternative IMF real exchange rate definition as objective variable.

4.9 Conclusions

This chapter aimed at studying the impact of leading commodity prices long run term trends and their volatility on the real exchange rate short term convergence in an error correction background in a selection of advanced and emerging countries, chosen according to literature based factors such as their leading commodity export sector share over the total export value. Through various specifications of transition regressions, our paper takes into account different measures of regime switch volatility in order to infer on the existence of arbitrage opportunities and the alternating regimes of convergence of the exchange rate to its equilibrium, proving that the commodity points theory of Heckscher can effectively be generalized at the longitudinal level under specific assumptions related to country selection. In synthesis, our threshold analysis allowed us to infer whether or not the absolute size of the deviation (the exchange rate variation itself), the relative behavioral misalignment (the error correction term) or commodity point-led arbitrage (the commodity price index) matter or not across a selection of groups of advanced and non advanced countries. From our final estimates, divergence from the estimated equilibria appear to matter more when compared to absolute deviations of the exchange rate. The absolute value of the exchange rate (the fact of it being under or overvaluated) appears to be a possible leading variable only in the case of the seminal selection by Cashin, Céspedes, and Sahay (2004) (*group I*) or once a limit threshold of 20 percent for non-energy commodity exporters is considered. Finally, once international commodity prices are chosen as the threshold variable, "literature driven" discretional export thresholds which would make the relationship between exchange rates and commodity prices stable and significant in emerging and not emerging countries no longer apply, allowing commodity threshold theory to stand alone. Further studies related to the commodity-prices exchange rates nexus would as such need to find an alternative exogenous factor which would make the difference between specialized exporters and non specialized exporters more evident. As a final remark, we take notice of how in the overall panel computations the slope of the transitions tend to a large number, making the transition quasi-instantaneous.

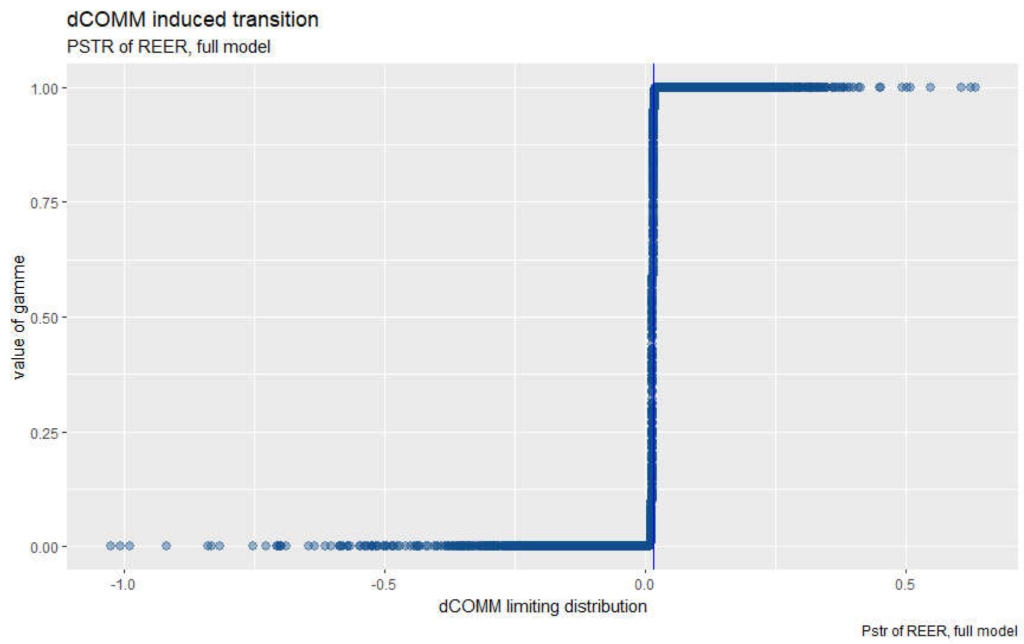

Figure 4-1: Transition function. Threshold variable: commodity price index variation.

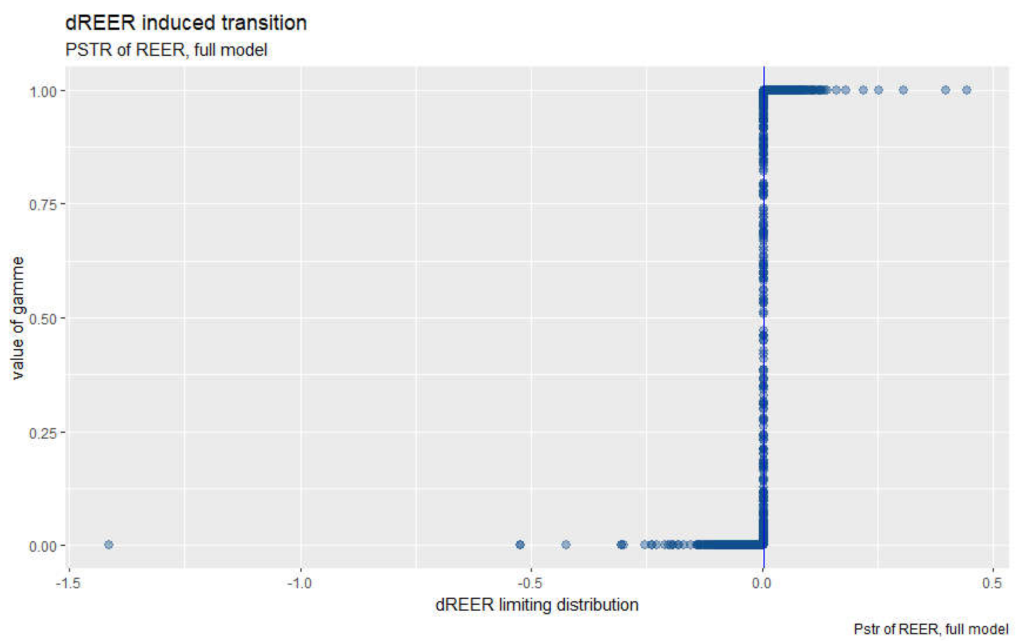

Figure 4-2: Transition function: Threshold Variable: exchange rate variation.

Chapter 5

Conclusions

5.1 Main findings

This thesis aimed at focusing on the behavior of the exchange rate as dictated by its behavioral fundamentals, with a specific focus on commodity prices. As we did that, we tried to focus, given the restriction of the limited temporal dimension of our analysis, on possible nonlinearities. We argued that the weakly exogenous relationship between exchange rates and commodity prices might have changed over time when compared to the crucial findings of Cashin, Céspedes, and Sahay (2004), reflecting changes in the structure of the economy, but concluded that nevertheless such relationship is still present, although with the caveats illustrated in the main body of our thesis, for commodity exporting emerging countries.

5.1.1 Chapter two

In chapter two, our initial analysis proves that, at the country case level, price differentials behavior of some small commodity exporters can be represented with satisfying precision by threshold models which have found their foothold in the literature of the nineties and the two thousands. As we run an out-of-sample forecast exercise, we showed how the notion of threshold convergence of price differentials to their unconditional average is however overshadowed by alternative non-linear models unrelated to literature. The idea of linear purchasing power parity appears to still be valid in some countries, so no relevant different speeds of convergence are detected in some of our country results, but a threshold representation could still be confirmed in a handful of them, making us wonder if a generalization of the Heckscher commodity points theory would still be possible.

5.1.2 Chapter three

As we moved on to chapter three looking for a proof of generality of either the linear or threshold convergence, we underwent a panel analysis taking into account two potential additional sources of nonlinearity: cross section heterogeneity and global shocks. Building an ad-hoc commodity price index with new updated trade weights, our estimate showed lower elasticities in the long run with respect to those estimated in the past, and a conclusive panel Granger causality testing allowed us to conclude that the (weak) causation we were looking for in emerging exporting countries and non energy exporting countries appear to be verifiable only in the short term, in the Granger sense.

5.1.3 Chapter four

Finally, in chapter four, we choose to merge the panel framework with the threshold modelling from the second chapter, trying to check whether or not an ideal exogenous source of variation like our newly built international price index could contribute in deciding the varying speed of adjustment of the exchange rate. As we could confirm that the absolute source of variation and the misalignment of the exchange rate from its behavioral equilibrium appear to be cofactors into explaining threshold convergence, we could find decisive evidence that the commodity price index, which would ideally embody the commodity points theory of Heckscher, can effectively be generalized at the longitudinal level, but its effect do not seem to be entirely conditional on the discretional restriction related to the weight of the commodity trade over the total trade of any of our country groupings. The absolute value of the exchange rate itself appears to perform rather well in describing the alternate speed of transition, which appears much steeper than what has been found out to be in previous country studies.

5.2 Policy making implications

The degree of dependence of emerging economies to exported commodity price fluctuations brings afloat, in terms of policy making indications, the possibility of anchoring the nominal value of the exchange rate to the price of the most exported commodities, in the spirit of Frankel and Saiki (2002b) and Frankel (2005b). However, in contrast to what these economist have proposed in order to counteract exchange rates shocks, our research leads us to conclude that such a possibility could never be advisable as a generalized policy for commodity exporting emerging countries. This appears evident in chapter four, where we tested for possible deviations from parity led by commodity price volatility and verified its feasibility for a number of country groupings but not for those belonging to the seminal Cashin, Céspedes, and Sahay (2004) selection, where commodity exporting countries where chosen with a cut off percentage of fifty percent of commodity trade over total trade. In a sense, the only way to suggest a policy is to accept the logical idea that it needs to be as discretional as it is needed: in chapter two, although with a very bad out-of-sample performance, we have seen autoregressive threshold modelling perform "decently" for a limited number of countries. Those are the countries to whom we would suggest to follow the advice to anchor the nominal exchange rate to the variation of either its most exported commodity or a basket of its most exported commodities.

5.3 Future developments

In this section of the thesis we argue on possible future lines of research related to the contents of previous chapters.

5.3.1 Chapter two

In chapter two, we saw how threshold models would not be the most suitable model in terms of forecastability for out-of-sample estimates of the exchange rate behavior. On the other side, additional nonlinear models like the additive model and the simple one layer neural network we employed gave forecasting results at least comparable to the benchmark autoregressive model. Future research should then focus on exploring, in a country-case study similar to ours, additional more complex forms of nonlinearity, which do not fall into any specific theory we have mentioned in our paper. In this regard, we reckon that adding to the complexity of the neural network model we employed would represent a good starting point, in conjunction with additional state models such as those depending on Markov-Switching theory.

5.3.2 Chapter three

In chapter three, we focused on how accounting for heterogeneity and cross section dependence would modify the elasticity of the exchange rate with respect to commodity price movements. Among the models we have employed, we reckon we would need an additional analysis which would merge the dynamic estimates of the Pedroni's DOLS method we employed together with the possibility of accounting for cross section dependence. This would further reduce the bias in our estimates. Furthermore, as we employed the panel Granger causality testing by Dumitrescu and Hurlin (2012), we might want to point out that a future version of our work in such chapter would have to evaluate and correct the test according to the degree of heterogeneity of our cross sections. Indeed, the test from Dumitrescu and Hurlin (2012) appears to have enough power only when the alternative hypothesis is very homogeneous.

5.3.3 Chapter four

In chapter four finally, we have seen how commodity prices would represent an alternative threshold to look at how the exchange rate transitions from one state to another. However, given our results, and considering the fact that commodity prices do not happen to be a reliable threshold for emerging exporting countries with a very high export share, we ought to keep searching for an exogenous, non derived measure[1] that would dominate the transition in every single possible country group we consider.

Above all, we should point out that our analysis does end in 2012 and it is motivated by the existence of a newly derived commodity price index whose variation is object of study in the thesis. As time progresses, much attention will have to be put into gathering

[1] That would be a measure which is neither a result of a combination of its behavioral elements (the real exchange rate variation), neither its misalignment.

series to add to the length of the analysis in order to get more decisive evidence over the validity of exchange rate convergence theories.

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