The Changing Relationship Between Commodity Prices and Prices of Other Assets with Global Market Integration

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Abstract

We explore the linkage between equity and commodity markets, focusing in particular on its evolution over time. An important debate in the literature concerns whether the large fluctuations in commodity prices in the late 2000s can be attributed to less segmentation of commodity markets. We document that a country's equity market value has significant out-of-sample predictive ability for the future global commodity price index for several primary commodity-exporting countries. We find, however, little evidence of in-sample predictive ability, even after allowing for instabilities. The out-of-sample predictive ability of the equity market appears towards the middle of the 2000s, thus suggesting a decrease in market segmentation and the possibility that shocks in equity markets might have started to spill-over onto commodity markets in that period. The results are robust to using country-specific commodity indices as well as firm-level equity data. Finally, our results indicate that exchange rates are a better predictor of commodity prices than equity markets. We show, however, that this is not the case for non-commodity currencies, as intuition would suggest.

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1 Introduction

In a recent paper, Chen, Rogoff and Rossi (2010) showed that exchange rates of small open economies with a large export share of primary commodities ("commodity currencies") may offer an alternative approach to forecasting commodity prices. In particular, Chen Rogoff and Rossi (2010) focus on Australia, Canada, Chile, New Zealand, and South Africa, which are a set of small commodity exporters with market-based floating exchange rates. They show that the exchange rates of such countries vis-a'-vis the U.S. Dollar have predictive content for future commodity price indices, either country-specific indices or the IMF aggregate commodity price index. They justify their empirical findings by using a present value argument: exchange rates, like any asset prices, should be determined as the net present value of fundamentals. Chen, Rogoff and Rossi (2010) identify commodity prices as a fundamental for small open economies whose exports heavily rely on primary commodities. They argue that the present value model thus implies that commodity prices should Granger-cause exchange rates, a finding that is empirically supported by the data, especially after controlling for instabilities. Given that commodity prices are "essentially exogenous" for small open economies, they interpret the Granger-causality finding as empirical support for the net present value theory of exchange rate determination.

The paper by Chen, Rogoff and Rossi (2010) establishes a structural link between exchange rates and future commodity prices through the terms of trade and income channel. Alternatively, one might conjecture a financial linkage across asset markets, where equity or bond markets in these countries also offer useful information for commodity market behavior. In an insightful discussion of Chen, Rogoff and Rossi (2010), Helene Rey (American Economic Association meetings, 2009) showed suggestive evidence that the Australian, Canadian, and Chilean stock price indices have predictive ability for commodity price indices, similar to that of the exchange rates. In this paper, we further explore the linkage between equity, commodity, and the exchange rate markets, focusing in particular on studying its evolution over time.

Our main findings are as follows. First, we document that a country's equity market value has significant out-of-sample predictive ability for the future global commodity price index for several commodity currencies. We find, however, little evidence of in-sample predictive ability, even after allowing for a one-time reversal in the performance of the models. The out-of-sample predictive ability of the equity market predictor appears towards the middle of the 2000s, thus suggesting a decrease in market segmentation and the possibility that shocks in equity markets might have started to spill-over onto commodity markets in that period.

We also compare the performance of equity markets as predictors for the global commodity price index with that of exchange rates. In particular, Chen et al. (2010) have found that exchange rates of commodity currencies are a reliable predictor for the commodity price index, and we investigate whether equity market values might provide further improvements relative to using exchange rates as predictors. Our results indicate that exchange rates are a better predictor than equity markets, at least in-sample. We show, however, that this is not the case for non-commodity currencies, as intuition would suggest.

We then turn to a more thorough analysis of the predictive ability of equity markets. In particular, one might expect that country-specific equity markets might be better predictors of country-specific commodity price indices, and we test that conjecture. We indeed find that country-specific equity markets do Granger-cause country-specific commodity price indices, especially after taking instabilities into account. We also find that they have out-of-sample forecasting ability as well. Interestingly, the results are robust to using firm-level equity prices.

Our paper is related to several strands of the literature. There is an increasing literature on the linkages between commodity prices and other asset prices/ markets to which our paper is related, among which the works by Tang and Xiong (2010) and Buyuksahin, Haigh and Robe (2008) are of particular interest. Tang and Xiong (2010) observe that increasing investment levels have flowed into commodity markets between 2006 and 2010, which spurred a debate on whether speculation might have caused excessive increases in the cost of primary commodities (including energy and food) and their volatility. There are two opposite views: one the one hand, Hamilton (2009) and Kilian (2009) have argued that the increase in commodity prices in that period was the consequence of the rapid growth in emerging economies, whose demand soared, pushing commodity prices up. On the other hand, the increase in commodity prices might have been caused by large flows of investments in commodity indices; in fact, the estimated investment in commodity indices (and related instruments) increased from $15 billion dollars in 2003 to $200 billion in mid-2008). Tang and Xiong (2010) note that a large literature pointed to the segmentation of commodity markets from other financial markets before 2000 (see e.g. Gorton and Rouwenhorst, 2006); however, there has been a sizeable increase in investment in commodity markets since then. As evidence of the tendency towards less segmentation, Tang and Xiong (2010) show that one-year rolling return correlations of crude oil with returns of other commodities (such as soybeans, cotton, live cattle, and copper) have increased dramatically since the early 2000s. They also argue that the rapid economic growth of emerging economies cannot be the main responsible to the soar in commodity prices since future prices of several commodities in China behaved very differently from those in the US. They also argue it cannot be caused by general inflation, since inflation in that period was mostly constant. Instead, they argue that the main responsible was the increasing presence of investors in the commodity markets, who generated a spill-over of shocks from outside commodity markets into the commodity markets. As additional support of their argument, they show a dramatic increase in the volatilities of commodity markets since 2004, which coincides with the peak of the increase in the volatilities in oil returns and the world equity index; furthermore, they show that the volatility in individual commodities was partially driven by the increased return correlation with oil, and that the indexed commodities experienced a higher increase in volatility. Their policy conclusions are that, although the increasing presence of commodity index investors seeking risk diversification might improve sharing of commodity price risk, and leading to improvements in the prices enjoyed by farmers and producers of commodities ("diversification benefit"), on the other hand they might also introduce more volatility from outside commodity markets into different commodities ("volatility spill-over effect"). The net effect is unclear, and thus the policy implications. The paper by Buyuksahin, Haigh and Robe (2008) reaches instead a very different conclusion: their correlation analysis shows a lack of greater return co-movement across equities and commodities, which suggests that commodity markets can still be used for portfolio diversification. These papers are closely related to ours: we also seek to study the relationship between commodity prices and equity markets, but we tackle the problem by looking at Granger-causality type of regressions (rather than correlations), which would be more informative regarding whether, for example, the behavior in equity markets has predictive ability for future commodity prices. We also pay particular attention to the existence of structural breaks, and provide a thorough analysis of the robustness of such relationship to instabilities as well as an indication of when the predictive ability was present. In our analysis, structural breaks can happen at an unknown point in time, which we can consistently estimate using the data; this is therefore another difference with the paper by Buyuksahin, Haigh and Robe (2008), who focus on given sub-samples in their analysis. Finally, our analysis applies to both global commodity price indices as well as country-specific price indices, whereas most of the literature focuses on global commodity price indices alone.

This paper is also more generally related to the literature on the predictive ability of commodity prices. In particular, there is a large literature on the relationship between the prices of primary commodities and exchange rates: Amano and Van Norden (1995), Chen and Rogoff (2003) and Cashin et al. (2004), among others. Alquist, Kilian and Vigfusson (2011) have recently provided a thorough overview of whether real or nominal oil prices are predictable based on macroeconomic aggregates. In particular, they review alternative specifications of the oil price variable used in empirical work (alternative oil price measures, choice of sample period, model specifications and alternative assumptions about future demand and supply conditions). They note problems with combining data from the pre-1973 and post-1973 period because of structural breaks, and emphasize the importance of discarding pre-1973 data. They find that nominal oil prices are predictable in-sample using lagged US CPI, money aggregates, global commodity prices, and exchange rates of commodity exporters. On the other hand, real oil prices are predictable in-sample using global real output. However, in-sample predictability need not translate in out-of-sample forecastability. In particular: (i) VARs provide better real oil price forecasts than a random walk benchmark up to 6 months, but at longer horizons the random walk is the best predictor;[1] (ii) growth of non-oil industrial raw material prices improve forecasts of nominal oil prices at horizons of 1-3 months; still, the random walk is best at longer horizons; (iii) futures do not improve forecasts. Finally, they demonstrate that there are only small gains in using oil prices to forecast real GDP growth, no matter whether one uses real or nominal oil price measures, or treats oil prices as exogenous or endogenous. They also find some evidence that non-linearities are important in forecasting: the 3-year net oil price increase is capable of improving forecasts of GDP growth. Finally, oil price volatility need not be an adequate measure of risk faced by market participants: the use of predictive densities is recommended. The main difference between Alquist et al. (2011) and this paper is that we focus on the predictive ability of commodity price indices, of which oil (and, more in general, energy) prices are only one component, and we consider equity market values as possible predictors.

The paper is organized as follows. First, we provide empirical evidence on the existence of a relationship between equity markets of commodity countries and the global commodity price index, paying particular attention to in-sample Granger-causality, instabilities and out-of-sample forecasting. We also compare the predictive ability of equity markets for the global commodity price index versus that of exchange rates, both in-sample and in out-of-sample forecast comparisons. Section 3 instead provides an analysis of whether equity markets are

[1] Results are robust to real-time data.

useful for predicting country-specific commodity price indices. We also validate our results using firm-level data on the stock value of a prominent Canadian oil producing and exporting firm. Section 4 concludes.

2 Equity Markets and Global Commodity Prices

We collect quarterly data on equity indices for a variety of commodity currencies. We focus on the same countries studied in Chen, Rogoff and Rossi (2010), namely Australia, Canada, Chile, New Zealand and South Africa. Datastream provides price index data on the market index for these countries. Table 1 provides details on mnemonics as well as starting date of the sample. We focus on data before the financial crisis since the global financial turmoil would obscure the relationships we are seeking to uncover. Thus, all the data end in 2008Q1.

Table 1. Equity Market Data Description

Country	Mnemonics	Starting Date
Commodity Currencies:		
Australia	TOTMAU	Jan 1973
Canada	TOTMCN	Jan 1973
Chile	TOTMCL	July 1989
New Zealand	TOTMNZ	Jan 1988
South Africa	TOTMSA	Jan 1973
Other Currencies:		
Germany	TOTMKBD	Jan 1973
Japan	TOTMKJP	Jan 1973
UK	TOTMKUK	Jan 1973
France	TOTMKFR	Jan 1973

Note. Data are from Datastream.

In addition, we also collect data on country-specific commodity price indices and the IMF aggregate (global) commodity price index, which is a world export-earnings-weighted price index for more than 40 primary products;[2] the data are similar to those in Chen et al. (2010), to which we refer for more details. In order to verify that the main results hold for the commodity currencies that we consider, but do not hold for currencies of countries which

[2] Here we focus on the non-oil commodity price index due to its longer span of available data.

are not heavy exporters of primary commodities, we also collect data on equity indices for other countries, in particular, Germany, Japan, U.K. and France.[3]

2.1 In-sample Predictive Content

If we were able to establish that equity market indices were useful predictors for the global commodity price index, this would have obvious policy implications: Central Banks and the IMF could use stock market data to help in the difficult task of forecasting commodity prices. Let cp_t^W denote the global commodity price index and m_t denote the equity market index of countries that are heavy exporters of primary commodities, such as Australia (denoted "AU"), New Zealand (denoted "NZ"), Canada (denoted "CA"), Chile (denoted "CHI") and South Africa (denoted "SA").

We first consider whether stock prices have in-sample predictive content for future values of the global commodity price index. We consider the following regression, where equity markets are used to predict the global commodity price index:

$$E_t \Delta cp_{t+1}^W = \beta_0 + \beta_1 \Delta m_t + \beta_2 \Delta cp_t^W, \quad t = 1, 2, ..., T. \tag{1}$$

We start by considering traditional Granger-causality (GC) tests. In particular, we test whether equity markets Granger-cause future commodity prices.[4] The p-values of the Granger-causality test are reported in Panel A in Table 2, in the row labeled "m_t GC cp_{t+1}^W". It is clear that there is little evidence of in-sample predictive ability. There is however some evidence that the commodity price index Granger-causes future equity market values (see the row labeled "cp_t^W GC m_{t+1}").

However, it is well-known that Granger-causality tests fail in the presence of instabilities – see Rossi (2005). Given the widespread empirical evidence of instabilities in commodity prices (e.g. Chen et al., 2010, Tan and Xiong, 2010), one might worry that the Granger-causality tests reported in Table 2 may suffer from instabilities. We therefore test whether the Granger-causality relationship is unstable over time using the QLR test (Andrews, 1993). Panel B in Table 2 reports p-values of the QLR test. In several cases, the p-values are close to or smaller than 5%, thus signaling the presence of instabilities.

Given the concerns about instabilities, we proceed to test for in-sample predictive ability using Rossi's (2005) Granger-causality test robust to instabilities. Even after allowing for

[3] We do not consider the US since commodity prices are typically quoted in US dollars.

[4] All tests are implemented with a heteroskedasticity and autocorrelation robust variance estimation (Newey and West, 1987) using a bandwidth equal to $T^{1/3}$.

instabilities, there is little empirical evidence that equity market values Granger-causality the global commodity price index – see Panel C in Table 2 (row labeled "m_t GC cp^W_{t+1}"), although strong evidence that the commodity price index Granger-causes future equity market values (see Panel C, row labeled "cp^W_t GC m_{t+1}").

Table 2. Global Commodity Price Index: Granger-causality Analysis

	AUS	NZ	CA	CHI	SA
Panel A. Granger-Causality Tests					
m_t GC cp^W_{t+1}	0.509	0.361	0.584	0.303	0.563
cp^W_t GC m_{t+1}	0.015	0.376	0.031	0	0
Panel B. Andrews' (1993) QLR Test for Instabilities					
m_t GC cp^W_{t+1}	0.036	0.055	0.053	0.010	0.074
cp^W_t GC m_{t+1}	0.069	0.309	1	0	1
Panel C. Granger-Causality Tests Robust to Instabilities, Rossi (2005b)					
m_t GC cp^W_{t+1}	0.358	0.215	0.286	0.102	0.483
cp^W_t GC m_{t+1}	0	0.489	0	0	0

Note. Panels A-C report p-values for tests for $\beta_0 = \beta_1 = 0$ based on two regressions: (i) $\Delta cp_{t+1} = \beta_0 + \beta_1 \Delta m_t + \beta_2 \Delta cp_t$ (labeled "m_t GC cp_{t+1}") and (ii) $\Delta m_{t+1} = \beta_0 + \beta_1 \Delta cp_t + \beta_2 \Delta m_t$ (labeled "cp_t GC m_{t+1}"). Asterisks indicate significance levels at 1% (***), 5% (**), and 10% (*), respectively.

2.2 Out-of-sample Forecasting Ability

Granger-causality results are useful tools to analyze historical data; however, policy-makers as well as institutions such as the IMF would find useful to assess the existence of the predictive ability in real-time. In order to evaluate the real-time out-of-sample forecasting ability of the model, we produce a sequence of rolling out-of-sample forecasts based on the model with equity market value, eq. (1), and compare it with forecasts based on two benchmark models: the autoregressive (AR(1)) model:

$$E_t \Delta cp_{t+1} = \gamma_{0t} + \gamma_{1t} \Delta cp_t,$$

and the random walk (RW) model:

$$E_t \Delta cp_{t+1} = 0.$$

Panel (a) in Table 3 reports the Clark and McCracken (2001) test for equal predictive ability of the model (1) and the AR(1) benchmark whereas panel (b) reports the same test for comparing the Mean Squared Forecast Error (MSFE) of the model (1) with that of the random walk. Asterisks denote significance: at 1% (***), 5% (**), and 10% (*) respectively. We find that, for several countries, the model with equity market forecasts significantly better than the benchmarks, especially the random walk benchmark. There is less evidence of out-of-sample predictive ability of the commodity price index for future equity market values.

Table 3. Tests for Out-of-Sample Forecasting Ability

AUS	NZ	CAN	CHI	SA
Panel (a): Autoregressive benchmark				
A. MSFE differences: Model: $E_t\Delta cp_{t+1} = \beta_{0t} + \beta_{1t}\Delta cp_t + \beta_{2t}\Delta m_t$ vs. AR(1): $E_t\Delta cp_{t+1} = \gamma_{0t} + \gamma_{1t}\Delta cp_t$				
0.74**	0.91*	1.53	0.44	1.55
B. MSFE differences: Model: $E_t\Delta s_{t+1} = \beta_{0t} + \beta_{1t}\Delta s_t + \beta_{2t}\Delta cp_t$ vs. AR(1): $E_t\Delta s_{t+1} = \gamma_{0t} + \gamma_{1t}\Delta s_t$				
0.68	0.40	2.84	0.97	1.67
Panel (b): Random walk benchmark				
A. MSFE differences: Model: $E_t\Delta cp_{t+1} = \beta_{0t} + \beta_{1t}\Delta m_t$ vs. Random walk: $E_t\Delta cp_{t+1} = 0$				
-0.722*	-0.814**	-0.21*	-0.42**	0.70*
B. MSFE differences: Model: $E_t\Delta m_{t+1} = \beta_{0t} + \beta_{1t}\Delta cp_t$ vs. Random walk: $E_t\Delta m_{t+1} = 0$				
-0.52***	-0.42**	-0.24	-0.31***	0.10

Note. The table reports re-scaled MSFE differences between the model and the benchmark forecasts. Negative values imply that the model forecasts better than the benchmark. Asterisks denote rejections of the null hypothesis that the random walk is better in favor of the alternative hypothesis that the fundamental-based model is better at 1% (***), 5% (**), and 10% (*) significance levels, respectively, using Clark and McCracken's (2001) critical values.

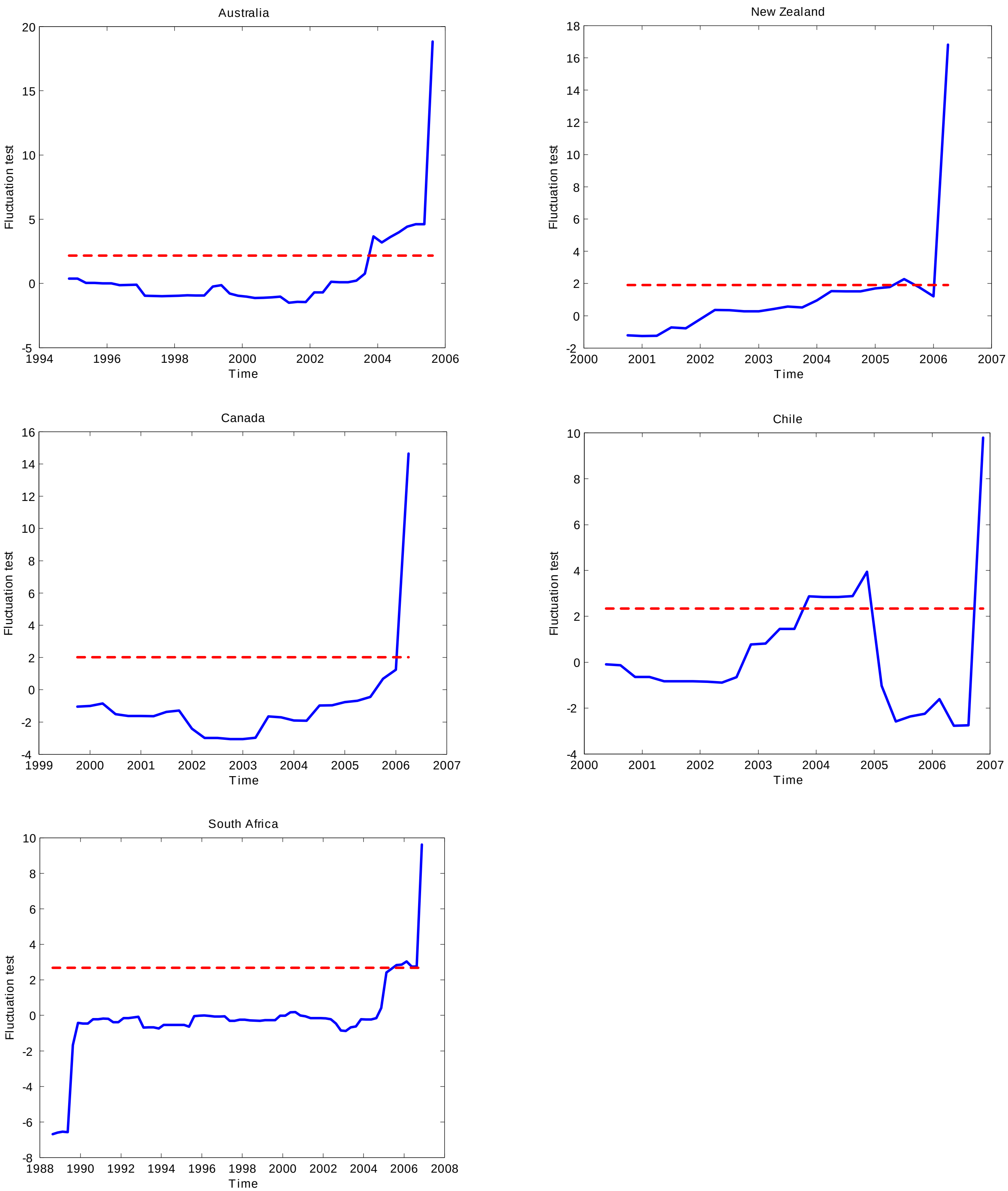
2.3 Forecasting Ability and Instabilities

It is important to note that the presence of instabilities (highlighted by the results in Panel B in Table 2) might also invalidate standard tests for forecast comparisons (see Giacomini and Rossi, 2010). This is particularly important for the analysis in this section. In fact, Tang and Xiong (2010) argue that the main responsible for the soar in commodity prices is speculation due to the increasing presence of investors in commodity markets, who may generate a spill-over of shocks from outside commodity markets into the commodity markets themselves. To empirically support their argument, Tang and Xiong (2010) report dramatic increases in the volatilities of commodity markets since 2004, which coincides with the increase in the volatility of oil returns and the world equity index. Alternative explanations involve increasing world demand for commodities, in particular from less developed countries, such as China. Thus, instabilities in the relationship between equity markets and commodity prices may provide useful insights on this important question: if the predictive ability of equity markets appears at the same time as the abrupt increase in investment in commodity indices, it might suggest that commodity markets have become less segmented around the same time.

Our objective is to provide additional empirical evidence on this important question by reporting forecast comparisons of model (1) against the random walk benchmark over time using the Fluctuation test developed by Giacomini and Rossi (2010). Giacomini and Rossi (2010) propose to measure of the local relative forecasting performance of the models, and test whether the competing models are equally good at forecasting the target variable at each point in time by plotting the (standardized) sample path of the relative measure of local performance together with critical values. The measure of local performance is obtained by the Clark and West (2007) test, which is appropriate for the nested models that we consider here. When the Fluctuation test is above the critical value, it signals that the model with equity information outperforms the autoregressive competitor at some point in time. The plot of the Fluctuation test provides some information on when the predictability appeared or disappeared over time, and thus it is particularly well-suited for our analysis. In fact, if the two models had similar and indistinguishable predictive ability up to mid-2000s and then afterwards the model with equity forecasted significantly better, then this might be an indication that commodity markets had become less segmented, and that there might have been spill-overs of shocks from equity markets into commodity markets, as the index of the latter became predictable using information from equity markets.

Figure 1 shows that indeed that was the case in the data. For all countries, we note that equity market values became significant predictors sometime in the mid-2000s, thus corroborating Tang and Xiong's (2010) results.

Figure 1. Fluctuation Test on Equity Markets and Global CP Index



2.4 Equity Markets vs. Exchange Rates

To perform a more thorough comparison between the model with exchange rates and the model with commodity prices in the presence of multiple instabilities, we use the test proposed by Giacomini and Rossi (2011). Giacomini and Rossi (2011) are concerned about comparing the performance of competing non-nested and possibly mis-specified models when their relative performance might be time-varying. They suggest to compare models' performance in rolling windows over the sample: measuring the performance over rolling window allows them to follow their relative performance as it evolves over time.

We compare the following two models:

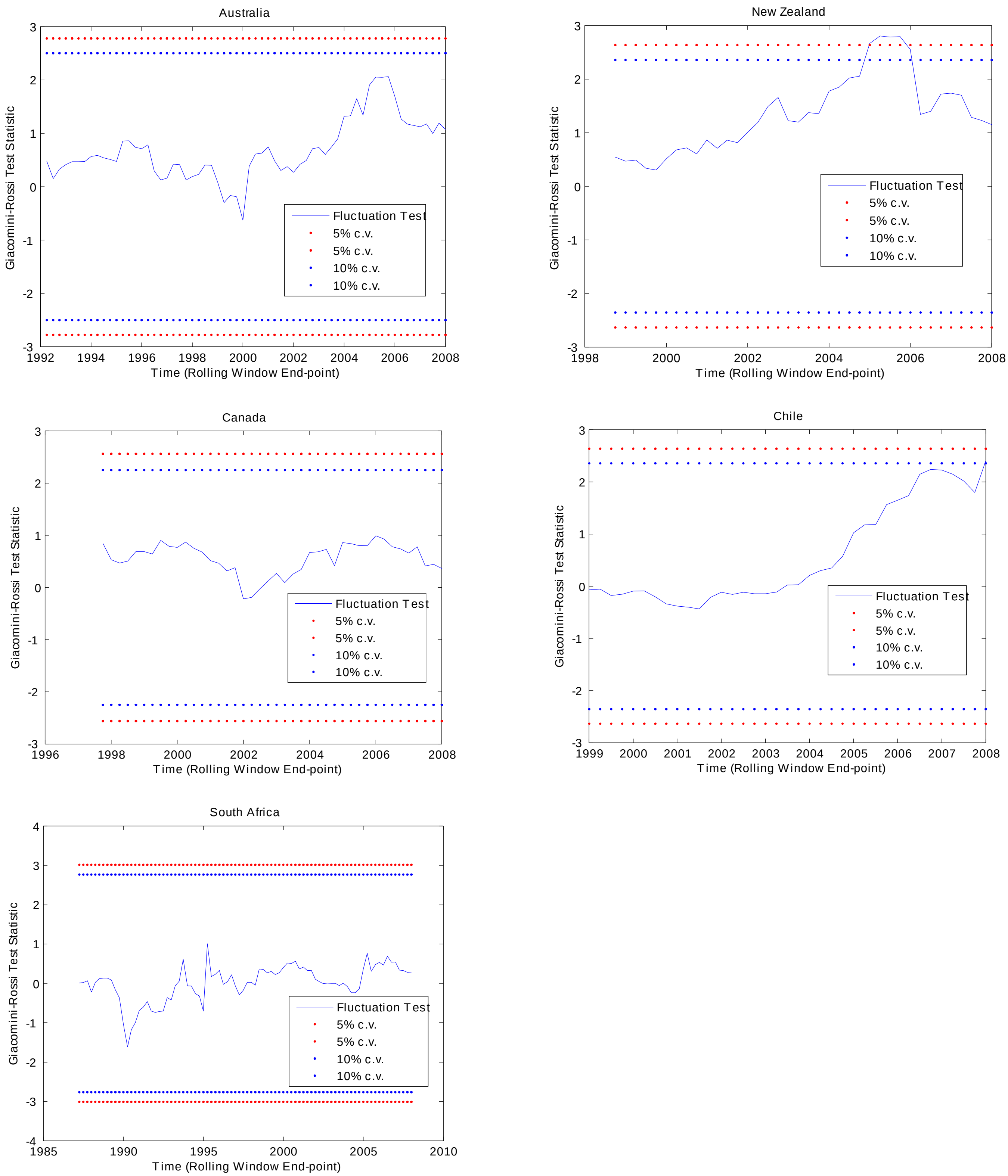
$$\Delta cp_{t+1}^{W} = \beta_0 + \beta_1 \Delta m_t + \beta_2 \Delta cp_t^{W}, \text{ and} \tag{2}$$

$$\Delta cp_{t+1}^{W} = \beta_0 + \beta_1 \Delta s_t + \beta_2 \Delta cp_t^{W}, \tag{3}$$

where Δm_t is the country-specific rate of growth of the equity market and Δs_t is the country-specific rate of growth of its exchange rate vis-a'-vis the U.S. dollar. We implement Giacomini and Rossi's (2011) test by calculating the difference between the in-sample squared fitted errors of models (2) and (3) in rolling windows over the sample. The size of the rolling window varies depends on the available sample: we use a window of 48 observations for Australia, 41 for New Zealand, 70 for Canada, 37 for Chile and 28 for South Africa.

Figure 2 reports the results. The figure reports, for each country, the test statistic (solid line) together with two-sided critical values at the 5% as well as the 10% significance levels, depicted in light and dark dots, respectively. When the test statistic is below the lower critical value line, we conclude that model using equity value performs the best; when the test statistic is above the upper critical value line, we conclude that the model using exchange rates performs significantly better; when the test statistic is always in between the critical value lines, that means that the two models are indistinguishable. Figure 2 shows that for Australia, Canada and South Africa, the fit of the two models is similar; however, in the case of New Zealand and Chile, using exchange rates to predict the global commodity price index provides a better fit to the data – at the 5% significance level for New Zealand and at the 10% significance level for Chile. Note that for all countries except South Africa, the test statistic is typically positive, pointing towards a better fit of the model with exchange rates even if it is not always significant in the data.

Figure 2. Giacomini and Rossi's (2010) Test: Equity vs. Exchange Rates

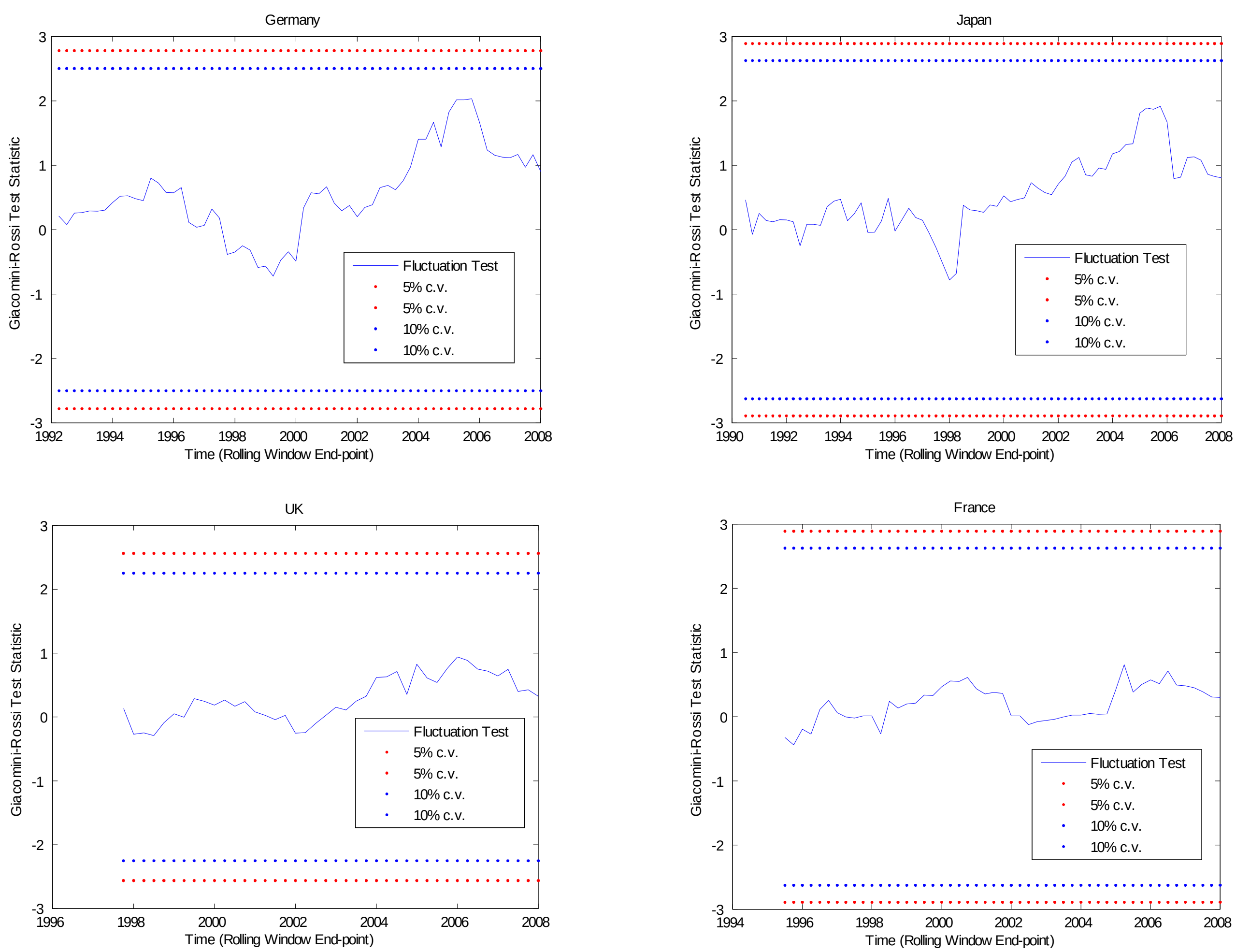


2.5 Robustness to Non-Commodity Currencies

One may also be interested in evaluating whether country-specific equity markets of countries that are not heavy exporters of commodities have predictive power for the world commodity price index; if that is the case, then the channel of transmission for equities is likely not related to the fact that such countries are heavy exporters of commodities but, rather, to other factors. We consider four such countries: Germany, Japan, UK and France.

Figure 3 reports the results of the Giacomini and Rossi's (2010) Fluctuation test. For none of these countries exchange rates are better predictors than equity markets, as expected.

Figure 3. Giacomini and Rossi's (2010) Test: Equity vs. Exchange Rates



2.6 Robustness to Out-of-sample Forecast Comparisons

Finally, we compare the forecasting ability of model (2) with that of model (3) out-of-sample. We use the Diebold and Mariano's (1995) test since the models are non-nested. The results, reported in Table 4, show that the two models are not statistically different from each other.

Table 4. Tests for Out-of-Sample Forecasting Ability

AUS	NZ	CAN	CHI	SA
Diebold and Mariano's test: MSE of Model: $E_t\Delta cp^W_{t+1} = \beta_{0t} + \beta_{1t}\Delta cp^W_t + \beta_{2t}\Delta s_t$ minus MSE of Model: $E_t\Delta cp^W_{t+1} = \beta_{0t} + \beta_{1t}\Delta cp^W_t + \beta_{2t}\Delta m_t$				
-0.062	0.855	1.119	0.623	0.881

Note. The table reports re-scaled MSFE differences between the model and the benchmark forecasts. Negative values imply that the model forecasts better than the benchmark. Asterisks denote rejections of the null hypothesis that random walk is better in favor of the alternative hypothesis that the fundamental-based model is better at 1% (***), 5% (**), and 10% (*) significance levels, respectively, using Clark and McCracken's (2001) critical values.

3 Equity Markets and Country-specific Commodity Price Indices

The analysis in the previous section demonstrated that equity markets have some out-of-sample predictive ability for the global commodity price index, and that such predictive ability started to show up in the data around mid-2000s. One might argue that country-specific equity market values might have more predictive ability for country-specific commodity price indices. We evaluate such conjecture in this section. To preview our results, we find strong in-sample Granger-causality of equity markets as well as stronger out-of-sample forecasting power. We show that the results are also robust to considering firm-specific equity values.

3.1 In-sample Predictive Content

We first consider whether stock prices have in-sample predictive content for future commodity prices. Let cp_t denote the country-specific commodity price index and m_t denote the country's equity market index. Consider the following regressions:

$$E_t\Delta cp_{t+1} = \beta_0 + \beta_1\Delta m_t + \beta_2\Delta cp_t, t = 1, 2, ..., T. \quad (4)$$

We start by considering traditional Granger-causality tests.[5] The results are reported in Panel A in Table 5. In the cases of Australia and South Africa, we do find that the rate of growth of equity markets significantly Granger-causes commodity prices at the 5% significance level.

The table also reports the reverse effect, namely the predictive ability of commodity prices for the equity markets. The latter is estimated from the regression:

$$E_t \Delta m_{t+1} = \beta_0 + \beta_1 \Delta cp_t + \beta_2 \Delta m_t, \tag{5}$$

Interestingly, the reverse effect is strongly significant for all countries.

Table 5. Bivariate Granger-Causality Tests

	AUS	NZ	CAN	CHI	SA
A. P-values of $H_0 : \beta_0 = \beta_1 = 0$ in $\Delta cp_{t+1} = \beta_0 + \beta_1 \Delta m_t + \beta_2 \Delta cp_t$					
	.028**	.379	.109	.429	.032**
B. P-values of $H_0 : \beta_0 = \beta_1 = 0$ in $\Delta m_{t+1} = \beta_0 + \beta_1 \Delta cp_t + \beta_2 \Delta m_t$					
	0***	.016**	0***	0***	0***

Note: The table reports p-values for the Granger-causality test. Asterisks mark rejection at the1% (***), 5% (**), and 10% (*) significance levels respectively, indicating evidence of Granger-causality.

We further test whether the Granger-causality relationship is unstable over time using the QLR test (Andrews, 1993). Table 6 reports the results. From Panel A, it is clear that the relationship between past equity market values and commodity prices has been subject to significant structural breaks in several countries, including Australia, New Zealand, South Africa, and marginally for Canada and Chile. Similar results hold for the reverse regression.

[5] All tests are implemented with a heteroskedasticity and autocorrelation robust variance estimation, see Newey and West (1987) using a bandwidth equal to $T^{1/3}$.

Table 6. Andrews' (1993) QLR Test for Instabilities in Bivariate GC Tests

	AUS	NZ	CAN	CHI	SA
A. P-values for stability of $(\beta_{0t,}\beta_{1t})$ in: $\Delta cp_{t+1} = \beta_{0t} + \beta_{1t}\Delta s_t + \beta_2 \Delta cp_t$					
	.00***	.052*	.13	.156	.00***
	(2004:2)				(2005:4)
B. P-values for stability of $(\beta_{0t,}\beta_{1t})$ in: $\Delta s_{t+1} = \beta_{0t} + \beta_{1t}\Delta cp_t + \beta_2 \Delta s_t$					
	.00***	.405	.875	.00***	.00***
	(2004:2)			(2005:1)	(2005:4)

Note: The table reports p-values for Andrew's (1993) QLR test of parameter stability. Asterisks mark rejection at the 1% (***), 5% (**), and 10% (*) significance levels respectively, indicating evidence of instability. When the test rejects the null hypothesis of parameter stability, the estimated break-dates are reported in the parentheses.

The Rossi's (2005) Granger-causality test robust to instabilities, reported in Panel A in Table 7, shows that the test finds much stronger empirical evidence in favor of predictability of equity markets for commodity prices than the traditional Granger-causality test. The results in favor of commodity prices Granger-causing future equity markets for commodity countries remain very strong – see Panel B.

Table 7. Granger-Causality Tests Robust to Instabilities, Rossi (2005b)

	AUS	NZ	CAN	CHI	SA
A. P-values for $H_0 : \beta_t = \beta = 0$ in $\Delta cp_{t+1} = \beta_{0t} + \beta_{1t}\Delta m_t + \beta_2 \Delta cp_t$					
	.00**	.30	.00***	.19	.00***
B. P-values for $H_0 : \beta_t = \beta = 0$ in $\Delta m_{t+1} = \beta_{0t} + \beta_{1t}\Delta cp_t + \beta_2 \Delta m_t$					
	.00***	.18	.00***	.00***	.00***

Note: The table reports p-values for testing the null of no Granger-causality that are robust to parameter instabilities. Asterisks mark rejection at the 1% (***), 5% (**), and 10% (*) significance levels respectively, indicating evidence in favor of Granger-causality. $\beta_t = (\beta_{0,t}, \beta_{1,t})'$.

We also evaluate the real-time out-of-sample forecasting ability of the model. We find quite strong empirical evidence that the model with equity market forecasts better than either an autoregressive or a random walk benchmark – see Table 8.

Table 8. Tests for Out-of-Sample Forecasting Ability

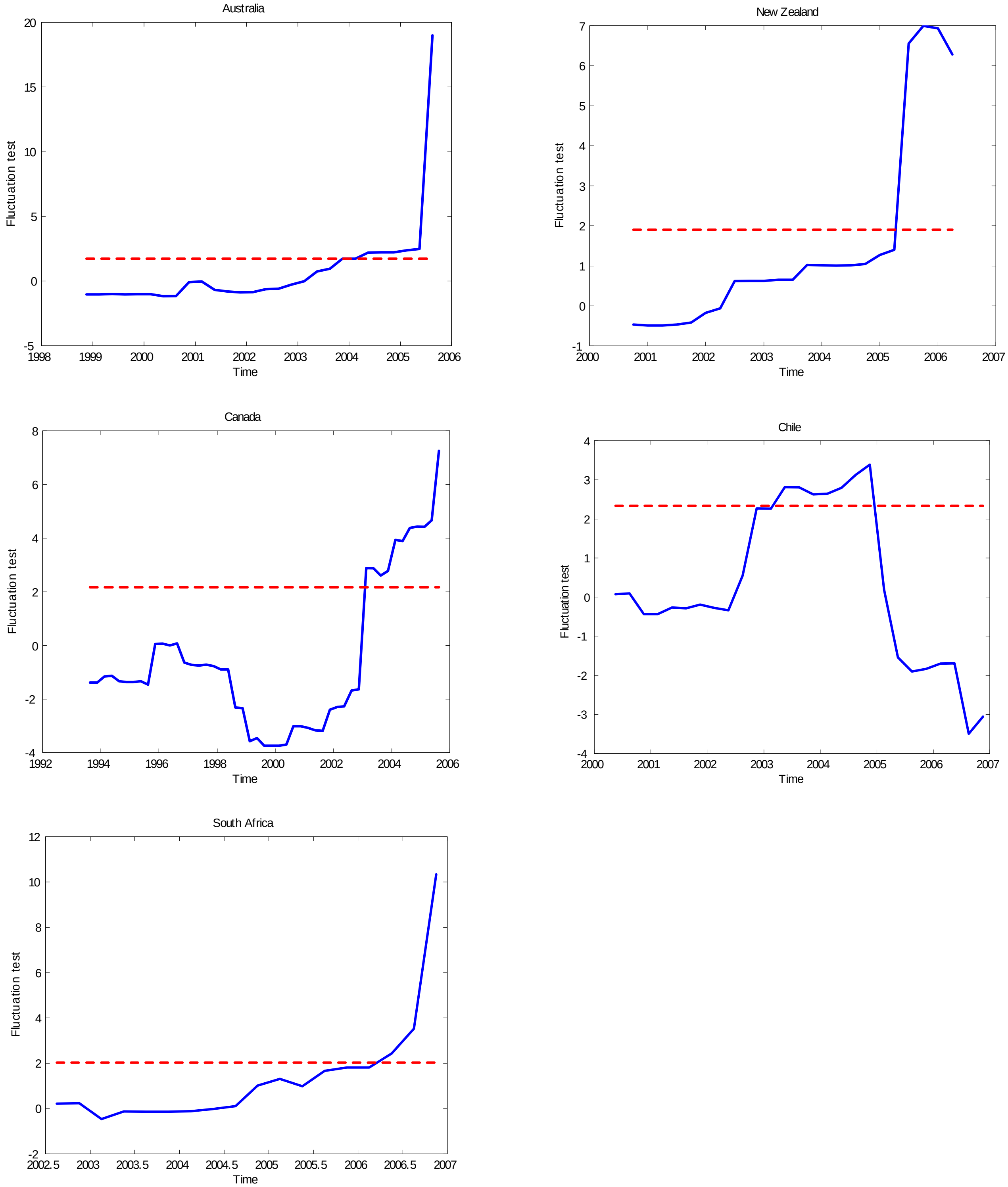
	AUS	NZ	CAN	CHI	SA
Panel (a): Autoregressive benchmark					
A. MSFE differences: Model: $E_t\Delta cp_{t+1}=\beta_{0t}+\beta_{1t}\Delta cp_t+\beta_{2t}\Delta m_t$ vs. AR(1): $E_t\Delta cp_{t+1}=\gamma_{0t}+\gamma_{1t}\Delta cp_t$					
	.520*	1.49***	1.58*	.24	.36***
B. MSFE differences: Model: $E_t\Delta s_{t+1}=\beta_{0t}+\beta_{1t}\Delta s_t+\beta_{2t}\Delta cp_t$ vs. AR(1): $E_t\Delta s_{t+1}=\gamma_{0t}+\gamma_{1t}\Delta s_t$					
	.59***	1.06	1.95	1.24	-.73***
Panel (b): Random walk benchmark					
A. MSFE differences: Model: $E_t\Delta cp_{t+1}=\beta_{0t}+\beta_{1t}\Delta m_t$ vs. Random walk: $E_t\Delta cp_{t+1}=0$					
	-1.59***	-1.82***	-.38	-.05	-1.83***
B. MSFE differences: Model: $E_t\Delta m_{t+1}=\beta_{0t}+\beta_{1t}\Delta cp_t$ vs. Random walk: $E_t\Delta m_{t+1}=0$					
	-1.18***	-.34	-.90	-.15***	-1.52***

Note. The table reports re-scaled MSFE differences between the model and the benchmark forecasts. Negative values imply that the model forecasts better than the benchmark. Asterisks denote rejections of the null hypothesis that random walk is better in favor of the alternative hypothesis that the fundamental-based model is better at 1% (***), 5% (**), and 10% (*) significance levels, respectively, using Clark and McCracken's (2001) critical values.

3.2 Instabilities in Forecasting Ability

As previously discussed, the presence of instabilities might also invalidate standard tests of forecast comparisons. We thus provide additional empirical evidence on the robustness of the predictive ability to the presence of instability by comparing the forecasts of model (4) against the random walk benchmark over time using the Fluctuation test developed by Giacomini and Rossi (2010). According to Figure 4, the predictive ability of the model with equity prices became much stronger towards the mid-2000s (only in the case of Chile, it became stronger towards the end of the sample but then disappeared again). These results strengthen the results we found when using equity markets to predict the global commodity price index: the predictive ability follows a pattern that is very similar across countries, notwithstanding the fact that the commodity price indices are different for each country, since they reflect the composition of their exports. We interpret the evidence as pointing towards to the synchronization of co-movements across commodity prices, as well as the co-movements between commodity prices and equity market values.

Figure 4. Fluctuation Test: Equity Market Predictors vs. Random Walk



3.3 Robustness to Firm-level Data

One might argue that our findings should be robust to using the equity market value of exporting firms. To verify whether the stock value of exporting firms has the same predictability properties, we collect data on firm level stock values. Unfortunately, several companies that export commodities are state-owned (e.g. the major industry in Chile producing copper) and several others are available only for a fraction of the sample we are considering. The results in this section use NASDAQ data from CRSP Monthly Stock database on the stock value of Imperial Oil Ltd., Canada's largest petroleum company. The company is engaged in the exploration, production and sale of crude oil and natural gas, and has been a leading member of the petroleum industry for more than a century, thus ensuring a long enough sample for empirical analysis.

Table 9 reports the results. Let f_t denote the firm's equity value at time t. We consider the following regression:

$$E_t \Delta cp_{t+1} = \beta_0 + \beta_1 (L) \Delta f_t + \beta_2 (L) \Delta cp_t. \tag{6}$$

Panel A shows that the stock market value of Imperial Oil Ltd. Granger causes the Canadian commodity price index, although the relationship is unstable over time (see Panel B). The Granger-causality is even stronger after allowing for instabilities (Panel C) and is robust in out-of-sample forecast comparisons (Panel D). Overall, our results are validated by firm-level stock price data.

Table 9. Canadian Commodity Price Index and Firm-level data

Panel A. Granger-Causality Tests

f_t GC cp_{t+1}		.01**
cp_t GC f_{t+1}		.75
Panel B. Andrews' (1993) QLR Test for Instabilities		
f_t GC cp_{t+1}		0.015** (2002:3)
cp_t GC f_{t+1}		0.523
Panel C. GC Tests Robust to Instabilities, Rossi (2005b)		
f_t GC cp_{t+1}		.00***
cp_t GC f_{t+1}		.87
Panel D. Out-of-Sample Forecasts		
	AR	RW
f_t forecast cp_{t+1}	-0.288***	-1.1433**
cp_t forecast f_{t+1}	1.082	1.9728

Note. Panels A-C report p-values for tests for $\beta_0 = \beta_1 = 0$ based on two regressions: (i) $\Delta cp_{t+1} = \beta_0 + \beta_1 \Delta f_t + \beta_2 \Delta cp_t$ (labeled f_t GC cp_{t+1}) and (ii) $\Delta f_{t+1} = \beta_0 + \beta_1 \Delta cp_t + \beta_2 \Delta f_t$ (labeled cp_t GC f_{t+1}). Estimated break-dates are reported in parentheses. Panel D reports results for out-of-sample forecast comparisons of the models against an autoregressive or a random walk. Asterisks indicate significance at 1% (***), 5% (**), and 10% (*) resp.

3.4 Equity Markets vs. Exchange Rates

To conclude, we compare the predictive ability of equity markets with that of exchange rates by considering multivariate regressions where country-specific commodity prices are predicted by both the country's equity market value and its exchange rate vis-a'-vis the US dollar. In particular, we augment eq. (4) with the rate of growth of the exchange rate, Δs_t, and estimate the following regression:

$$\Delta cp_{t+1} = \beta_0 + \beta_1 \Delta m_t + \beta_2 \Delta s_t + \beta_3 \Delta cp_t,$$

where Δm_t is the country-specific rate of growth of the equity market. Table 10 reports p-values of tests of Granger-causality: Panel A focuses on testing the significance of the equity market value and panel B focuses on testing the significance of the exchange rate. In both cases, there is little empirical evidence of Granger-causality. Similarly, there is little evidence that commodity prices help in predicting either exchange rates or equity market values.

Table 10. Trivariate Granger-Causality Tests

	AUS	NZ	CAN	CHI	SA
A. P-values of $H_0 : \beta_1 = 0$ in $\Delta cp_{t+1} = \beta_0 + \beta_1 \Delta m_t + \beta_2 \Delta s_t + \beta_3 \Delta cp_t$					
	.080*	.342	.702	.332	.986
B. P-values of $H_0 : \beta_1 = 0$ in $\Delta cp_{t+1} = \beta_0 + \beta_1 \Delta s_t + \beta_2 \Delta m_t + \beta_3 \Delta cp_t$					
	.977	0.128	.633	.034*	.301
C. P-values of $H_0 : \beta_1 = 0$ in $\Delta m_{t+1} = \beta_0 + \beta_1 \Delta cp_t + \beta_2 \Delta s_t + \beta_3 \Delta m_t$					
	0.66	.75	.99	.85	.09*
D. P-values of $H_0 : \beta_1 = 0$ in $\Delta s_{t+1} = \beta_0 + \beta_1 \Delta cp_t + \beta_2 \Delta m_t + \beta_3 \Delta s_t$					
	.170	.237	.701	.974	.703

Note: The table reports p-values for the Granger-causality test. Asterisks mark rejection at the1% (***), 5% (**), and 10% (*) significance levels respectively, indicating evidence of Granger-causality.

As Table 11 shows, however, most regressions are subject to instabilities. After controlling for such instabilities, Table 12 shows that the empirical evidence in favor of Granger-causality is stronger, no matter whether exchange rates or equity data are used.

Table 11. Andrews' (1993) QLR Test for Instabilities in Trivariate GC Tests

AUS	NZ	CAN	CHI	SA
A. P-values for stability of β_{1t} in: $\Delta cp_{t+1} = \beta_{0t} + \beta_{1t}\Delta m_t + \beta_2\Delta s_t + \beta_3\Delta cp_t$				
.00***	.587	0***	.446	.00***
(2004:2)				(2005:4)
B. P-values for stability of β_{1t} in: $\Delta cp_{t+1} = \beta_{0t} + \beta_{1t}\Delta s_t + \beta_2\Delta m_t + \beta_3\Delta cp_t$				
.741	.00***	.595	.787	.00***
	(2004:4)			(2005:4)
C. P-values for stability of β_{1t} in: $\Delta m_{t+1} = \beta_{0t} + \beta_{1t}\Delta cp_t + \beta_2\Delta m_t + \beta_3\Delta s_t$				
.00***	.02**	.11	.83	.00***
(2004:2)	(2004:3)			(2005:4)
D. P-values for stability of β_{1t} in: $\Delta cp_{t+1} = \beta_{0t} + \beta_{1t}\Delta s_t + \beta_2\Delta m_t + \beta_3\Delta cp_t$				
.00***	.00***	.743	.58	.00***
(2004:2)	(2004:4)			(2005:4)

Note: The table reports p-values for Andrew's (1993) QLR test of parameter stability. Asterisks mark rejection at the 1% (***), 5% (**), and 10% (*) significance levels respectively, indicating evidence of instability. When the test rejects the null hypothesis of parameter stability, the estimated break-dates are reported in the parentheses.

Table 12. Granger-Causality Tests Robust to Instabilities in Trivariate Regressions – Rossi (2005b)

AUS	NZ	CAN	CHI	SA
A. P-values for $H_0 : \beta_{1,t} = \beta = 0$ in $\Delta cp_{t+1} = \beta_{0t} + \beta_{1t}\Delta m_t + \beta_2\Delta s_t + \beta_3\Delta cp_t$				
0.14	1	.00***	0.64	0.04**
B. P-values for $H_0 : \beta_{1,t} = \beta = 0$ in $\Delta cp_{t+1} = \beta_{0t} + \beta_{1t}\Delta s_t + \beta_2\Delta m_t + \beta_3\Delta cp_t$				
1	.00***	1	.20	.00***
C. P-values for $H_0 : \beta_{1,t} = \beta = 0$ in $\Delta m_{t+1} = \beta_{0t} + \beta_{1t}\Delta cp_t + \beta_2\Delta s_t + \beta_3\Delta m_t$				
.00***	.055*	.32	1	.00***
D. P-values for $H_0 : \beta_{1,t} = \beta = 0$ in $\Delta s_{t+1} = \beta_{0t} + \beta_{1t}\Delta cp_t + \beta_2\Delta m_t + \beta_3\Delta s_t$				
.00***	.00***	1	1	.12

Note: The table reports p-values for testing the null of no Granger-causality that are robust to parameter instabilities. Asterisks mark rejection at the 1% (***), 5% (**), and 10% (*) significance levels respectively, indicating evidence in favor of Granger-causality.

Table 13. Tests for Out-of-Sample Forecasting Ability

	AUS	NZ	CAN	CHI	SA
Panel (a): Models' comparisons					
Diebold and Mariano's test: MSE of Model: $E_t\Delta cp_{t+1}= \beta_{0t}+\beta_{1t}\Delta cp_t+\beta_{2t}\Delta s_t$ minus MSE of Model: $E_t\Delta cp_{t+1}= \beta_{0t}+\beta_{1t}\Delta cp_t+\beta_{2t}\Delta m_t$					
	0.109	0.273	0.396	-0.428	1.316
Panel (b): Random walk benchmark					
MSFE diff.: Model: $E_t\Delta cp_{t+1}= \beta_{0t}+\beta_{1t}\Delta s_t+\beta_{2t}\Delta m_t+\beta_{3t}\Delta cp_t$ vs. RW: $E_t\Delta cp_{t+1}= 0$					
	-1.59***	-1.82***	-0.38	-0.05	-1.83***
Panel (c): Autoregressive benchmark					
MSFE diff.: Model: $E_t\Delta cp_{t+1}= \beta_{0t}+\beta_{1t}\Delta s_t+\beta_{2t}\Delta m_t+\beta_{3t}\Delta cp_t$ vs. AR(1): $E_t\Delta cp_{t+1}= \gamma_{0t}+\gamma_{1t}\Delta cp_t$					
	0.708*	1.602***	1.792	0.294	0.946

Note. The table reports re-scaled MSFE differences between the model and the benchmark forecasts. Negative values imply that the model forecasts better than the benchmark. Asterisks denote rejections of the null hypothesis that random walk is better in favor of the alternative hypothesis that the fundamental-based model is better at 1% (***), 5% (**), and 10% (*) significance levels, using Diebold and Mariano's (1995) and Clark and McCracken's (2001) c.v. in Panels (a) and (b-c), resp.

Table 13 provides an analysis of the forecasting ability of several models for predicting commodity prices out-of-sample. Panel (a) compares the model with exchange rates with the model with equity prices. The two models are non-nested since they use different predictors. Thus, we compare their forecasting ability by using the Diebold and Mariano's (1995) test. The results show that the two models are not statistically different from each other. When we instead compare the model using both equity as well as exchange rates with either a random walk (panel b) or an autoregressive model benchmark (panel c), we find that the model that uses both equity and exchange rates forecasts significantly better than the benchmarks for several countries.

4 Conclusions

This paper explored the linkage between equity and commodity markets, focusing in particular on studying its evolution over time.

The main findings are as follows. First, we document that a country's equity market value has significant out-of-sample predictive ability for the future global commodity price index for several commodity currencies, although little in-sample predictive ability. More interestingly, the appearance of the out-of-sample predictive ability of the equity market predictor can be dated around the middle of the 2000s. Since the mid-2000s marked a large increase in investment in commodity markets, our empirical evidence suggests a decrease in market segmentation at approximately the same time and the possibility that shocks in equity markets might have started to spill-over onto commodity markets in that period.

We provide a series of robustness analyses, including: (i) comparisons of the performance of equity markets as predictors for the global commodity price index with that of exchange rates; (ii) using other data definitions, such as country-specific commodity price indices and firm-level equity market data. The results indicate that exchange rates are a better predictor than equity markets, at least in-sample, and that our main results are robust to using other data definitions.

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