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## Carry Trade Returns and Commodity Prices under Capital and Interest Rate Controls: Empirical Evidence from China

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This paper examines the relationship between returns of carry trade and prices of commodity based collateral assets (aluminium, copper and gold) for the classical carry trade pair: U.S. Dollar and Chinese Yuan. Given the nature of the time series employed, we consider the presence of structural breaks for the empirical analysis. The Autoregressive Distributed lag (ARDL) model suggests that in the long run adding copper and gold in the carry trade portfolio reduces the standard deviation. Furthermore, the short-run dynamics only exist between gold price and returns of carry trade. Our causality results reveal that multi-horizon causality testing does uncover important information with respect to the dynamic interaction among carry trade returns and different commodity prices. In particular, we find a causal chain through copper price and broken causal chains between prices of aluminium and copper to carry trade returns (transmitted via gold price). We also use a structural VAR model to disentangle the underlying causes of gold price shocks. We show that close to 60% of the variation in the real price of gold can be attributed to structural shocks in the currency market.

**Keywords:** Carry trade, Commodity prices, Structural breaks, multiple horizon causality, Structural VAR

**JEL Codes:** G15 E44 C32

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# 1 Introduction

Understanding the exchange rate movements and their co-movement with primary commodities such as gold and copper is an important issue for portfolio management, developing optimal trading strategies and forecasting many macroeconomic variables. Depending on the phase of business cycle, commodity prices tend to move in the direction, which could be opposite to exchange rates, bond and stock market prices. This might suggest a potential arbitrage opportunity such as carry trade which involves speculation on several markets. In fact, over the few past decades, carry trade strategies have been profitable with high Sharpe ratios (Burnside et al., 2011; Laborda et al., 2014). Currency carry trade is an investment strategy consisting of borrowing in low interest rate currencies (short position in the funding currency) and lending in high interest rate currencies (long position in the target currency). In addition, carry trade is not a risk free strategy because it is subject to crash risk and liquidity squeezing characterized by the reversal of currency values between high and low interest rate countries (Laborda et al., 2014, p. 53).

The theoretical foundation for carry trade is the uncovered interest rate parity (UIP) condition. Assuming risk neutrality and rationality of investors, if the UIP holds, an investor should not be able to profit from such a strategy because any difference between the nominal interest rates will be revoked by the corresponding changes in the exchange rates (i.e., appreciation of the low interest rate currency and depreciation of the high interest rate currency). This means that a risk-neutral investor should be indifferent between foreign and domestic investment alternatives. However, as Fama (1984) has shown in his seminal paper, theoretical UIP tends to fail empirically. This failure or so-called forward bias puzzle is typically manifested as appreciation of low interest currencies and predictability of the currency excess returns suggesting that foreign exchange markets might not be efficient (Laborda et al., 2014). The existence of carry trade and the failure of UIP represent one of the major long standing puzzles in international finance which are typically explained as the evidence of time varying risk premium, expectational failure or both (MacDonald and Nagayasu, 2015).

Following Gagnon and Chaboud (2007), carry trade can be executed by the outright borrowing in the low interest rate currencies as well as through currency forward and futures contracts on a margin. In this case, the strategy is implemented on the foreign exchange markets by “taking long positions in the currencies that are traded on forward discount (high interest rate currencies) and short positions in currencies which are traded on forward premium (low interest rate currencies)” (Shehadeh et al., 2016, p. 375). In both cases, carry trade strategies tend to be highly leveraged. The financing for the carry trade can be executed with or without collateral

(so-called naïve carry trade). Paper and physical commodities, stocks, bonds and other assets can be used as the collateral. [Zhang and Balding \(2015\)](#) have shown that using collateral is not that uncommon in carry trade strategies especially for economic agents from developing countries such as China. Financial Times<sup>1</sup> reported that almost 70 percent of copper imports have been used as collateral for carry trade in China. While economic agents from developing countries might favour commodities as collateral for carry trade deals, investors from developed countries such as Japan tend to use shares of local companies ([Ferreira-Filipe and Suominen, 2014](#)). The exact relationship between the carry trade returns and prices of assets used as collateral is still far from being well understood because commodity prices are not only affected by the demand and supply for commodities but also by credit shocks in countries receiving carry trade deals ([Roache and Rousset, 2015](#)). Therefore, it is possible that credit shocks and capital controls may directly impact commodity prices through demand for collateral assets. The question arises however, what is the relationship between the carry trade returns and the prices of the assets used as collateral? What is the causality pattern (important for investors to gain benefits from carry trade) between the prices of the collateral and the carry trade returns? Specifically, if the direction of commodity prices affect returns of carry trade, then speculators are able to reduce risks of their carry trade portfolios by reacting to the corresponding fluctuations in commodity prices. These are the questions that this paper seeks to answer using U.S. dollar (USD) and Chinese yuan (CNY) carry trade pair, aluminium, copper and gold prices (collateral assets) and time series methods. US was chosen for this analysis due to a long history of being funding country for carry trade deals. China was chosen for this analysis for the following reasons. First, this is the country which simultaneously imposes capital controls and interest rate controls directly affecting the profitability of the carry trade strategies and the need of using collateral to bypass capital controls. Second, over the sample period China pegged exchange rate to the USD and kept closed capital account through imposing capital controls, being one step away from Triffen's trilemma.<sup>2</sup> Empirical evidence shows that monetary authority's attempts of maintaining fixed exchange rate, closed capital account together with the sovereign monetary policy often serve as the invitation for the speculators to engage in carry trade once the market conditions are ready.

This paper makes the following contributions to the literature. First, we offer alternative explanation of the failure of the UIP. Using the Autoregressive Distributed Lag (ARDL) model

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<sup>1</sup> Refer the article "China's low rates sound death knell for copper carry trade" on the Financial Times website, available at: <https://www.ft.com/content/e0b01e0e-20bd-11e5-ab0f-6bb9974f25d0>

<sup>2</sup> In open economies, the three features that policy makers would prefer their monetary system to achieve are exchange rate stability, freedom of financial flows and monetary policy autonomy. Nevertheless, at most two can coexist. Therefore, the trilemma refers to the inconsistency in policy regime, among financial controls, fixed exchange rate and floating exchange rate. Each of the the policy regimes is consistent with the two goals.

with structural breaks, we investigate the long-run relationship effects between collateral prices and carry trade returns. This allows us understanding whether there are common factors that drive collateral prices and carry trade returns together over the long term.

Second, we study causality between the prices of collateral assets and carry trade returns. To the best of our knowledge, only a few studies allow for structural changes when testing for causality. The models adopted in the literature vary, with conventional models involving standard Granger causality type setting. However, such conventional models may be misspecified so that they are unable to uncover all possible causal relationships between commodity prices and carry trade returns. For instance, the macroeconomic impact of the carry trade unwinding on commodity prices might occur in a nonlinear way (see [Fry-McKibbin et al., 2016](#)) rather than the linear way implied from the linear vector autoregression (VAR) models widely used in the literature. Furthermore, causality effects can arise through conditional volatility or at different frequencies rather than the aggregated used in many research. An important issue pertains with the number of variables included in the causality framework. Most studies have used bivariate VAR to investigate the causal relationship between commodity prices and returns of carry trade. The bivariate approach has become very popular in analyzing Granger-causality relationships because one-step ahead causation indicates  $h$ -step ahead causation (direct causality) between the two variables of interest. Nevertheless, [Lütkepohl \(1982\)](#) argued the issue of omitted variable(s) related to the bivariate setting, which can lead to erroneous conclusions with respect to causality inference. Based on this fact, [Lütkepohl \(1982\)](#) emphasized using multivariate models to examine causality patterns between two variables of interest because higher dimensional time series models can provide additional information on multiple causal channels among the system variables, that may remain hidden or lead to spurious correlations in the bivariate framework. Thus, adopting multivariate models can help so that useful information is not omitted, which further allows for the presence of causal chains among the system variables. In particular, one-step ahead non-causation implies  $h$ -step ahead non-causation in the bivariate framework, however, this does not necessarily valid in a multivariate model in which more than two variables of interest are included ([Dufour and Renault, 1998](#); [Lütkepohl, 1993](#)). By contrast, additional variables can induce indirect causality results at higher forecast horizons, indicating nuanced details on multiple-horizon causation which would be collapsed out in a bivariate setting. More importantly, all existing research only examine the direct causality between commodity prices and carry trade returns. Therefore, although they may avoid the problem of omitted variable(s) by adding relevant variables in the model, they cannot capture all possible causal links (indirect causality) that can show up at higher forecast horizons.

To address the issues mentioned above, we employ the [Hill \(2007\)](#) efficient tests of long-run

causation in trivariate VAR systems. The method applies the approach proposed by [Dufour and Renault \(1998\)](#) which extends the original [Granger \(1969\)](#) causality definition and is based on linear predictability at higher forecast horizons. This approach is attractive because it reduces the increasing complexity of nonlinear non-causality parametric restrictions in VAR models to linear, for horizons greater than one (higher horizons). The test allows us to investigate the dynamic interaction between collateral prices and carry trade returns and also helps us to provide additional information on both the time profile of causal effects and their direct and indirect nature.

Third, in our analysis we consider structural change by testing for structural breaks. The issue of structural break is of considerable importance in the analysis of macroeconomic time series. Structural break appears for variety of reasons, including financial crises, changes in institutional arrangements, regime shifts and policy changes. An associated problem is that of testing the null of structural stability against the alternative of one or multiple structural breaks. More importantly, if such breaks are present in the data generating process, but not allowed for in the specification of an econometric model, results can be biased towards the non-rejection of a false unit root null hypothesis ([Perron, 1989](#); [Perron, 1997](#)). The economic implication of such a result is to erroneously conclude that the examining series has a stochastic trend. It is, therefore, essential to allow for structural change in the data so as to more reliably implement time series analysis.

Fourth, we set up a revised [Frankel and Rose \(2010\)](#) model, in order to investigate the factors affecting commodity prices under interest rate control regime. We disentangle the causes underlying gold price shocks. In particular, we model changes in the real price of gold as arising from two different sources: shocks to the futures price in the last period (futures contracts can be used as alternative for carry trade strategies) and shocks to the carry trade returns.

Our results show that copper price has a positive impact on the carry trade returns in the long term. Conversely, there is an inverse relationship between gold price and returns of carry trade. This may be due to the fact that copper and gold are used as collateral for implementing carry trade strategies more often as compared to aluminium. Moreover, the tiny positive effect of copper price and large negative effect of gold price suggest that in the long run there are hedge characteristics for copper and gold returns on carry trade returns. We do not find a short-run association between the prices of aluminium and copper and carry trade returns. In contrast, gold price in the short-term does affect carry trade returns. Our causality results indicate a causal chain through copper price and there are broken causal chains between prices of aluminium and copper to carry trade returns (transmitted via gold price). In addition, based on an identified SVAR, we find that shocks in the currency market account for about 60% of the long run fluctuations of the

real price of gold.

The remainder of the paper is structured as follows: [Section 2](#) provides literature review, [Section 3](#) describes the mechanic of using commodities as collateral, [Section 4](#) and [5](#) discuss the data and empirical methodology employed in this study respectively, [Section 6](#) presents results of the study, [Section 7](#) reports the robustness checks performed to cross-validate the results by examining the influence of the way to use financial derivatives on collateral prices, and [Section 8](#) concludes the paper. In the Appendix, we provide the cumulative sum (CUSUM) test for the ARDL model and rolling window trivariate VAR order selection.

## 2 Review of Literature

Financial literature and in particular traditional factor models (e.g., [Mark, 1988](#); [McCurdy and Morgan, 1991](#); [Bansal and Dahlquist, 2000](#)) for exchange rates suggest that exchange rates are related to equity and debt markets. As argued in [Lustig et al. \(2011, p. 530\)](#), “profitability of currency trading strategies depends on the cost of implementing them” and hence it depends also on the cost of financing, further the cost of the collateral. However, the relationship between the source of financing, credit, trade flows and collateral prices for the loan is not well understood. One explanation was recently proposed by [Ready et al. \(2017\)](#), based on the general equilibrium model of international commodity trade and currency pricing, they suggested that the currency carry trade returns are related to the patterns in international trade: countries that specialize in exporting basic goods such as raw commodities tend to exhibit high interest rates (e.g., Brazil and China), whereas countries primarily exporting finished goods have lower interest rates on average (e.g., Japan and U.S.). This finding implies that there is an inner connection between commodity and carry trade, that is, appearance of profitability of carry trade may because of the commodity specialization of each country.

Another gap in the literature is the impact of carry trade and in particular capital flows from the funding country on asset prices in investment country and prices for collateral asset/commodity in the funding country. This gap has recently been studied by [Plantin and Shin \(2007\)](#), [Agrippino and Rey \(2013\)](#), and more recently by [Zhang and Balding \(2015\)](#), but these studies did not analyze the impact of carry trade on the prices and volatilities of the collateral assets. [Agrippino and Rey \(2013\)](#) examined cross border flows and credit to the investment countries (Australia) and found that departures from UIP and profitability of the carry trade can be brought by several factors including feedback loop between the U.S. and Australia monetary policy, capital inflows and credit creation. Specifically, more credit inflows into Australia tend to be associated with an appreciating exchange rate. Based on the conventional Johansen cointegration

test and vector error-correction model (VECM), [Zhang and Balding \(2015\)](#) found a long-run relationship between the copper stock value and USD-CNY carry trade returns. Their results suggest that it takes on average three weeks for increase in the covered carry trade returns to have a short-run increase on the copper stock value. Furthermore, based on the Toda Yamamoto version of Granger causality test, they found evidence in support of a short-run causality running from carry trade returns to copper trade financing. Although [Zhang and Balding \(2015\)](#) investigated both long-run and short-run relationship between the commodity value and carry trade returns, the Toda Yamamoto test they used has a number of drawbacks including the issue of low power.

### 3 Mechanism of carry trade with commodities as collateral

The carry trade deals using collateral are flexible. On average, such deals follow the similar framework as described by [Zhang and Balding \(2015\)](#) and [Tang and Zhu \(2016\)](#) and summarized in [Fig.1](#). [Fig.1](#) demonstrates the causes and effects of financing carry trade through using commodities as collateral. The consumer of commodities imports commodities from the representative producer. Both exporting and importing countries have futures markets, but importing countries (e.g., Brazil and China) have capital controls. Because commodities are not regarded as capital flows, they are not affected by the capital control regime. If the importing countries have high unsecured interest rates as compared to the exporting countries, demand for using commodities as collateral substantially increases ([Tang and Zhu, 2016](#)). Financial investors in the importing countries borrow foreign currency at low unsecured interest rates (Step 1) and on the borrowed funds they purchase commodities (e.g., copper, gold or iron ore) (Step 2). Then commodities are imported into the country and used to obtain domestic secured low interest loan (Step 3). To hedge commodity price risk, financial investors in the importing country can use local futures market. Furthermore, to hedge currency risk, investors can trade currency forward on the foreign exchange market.

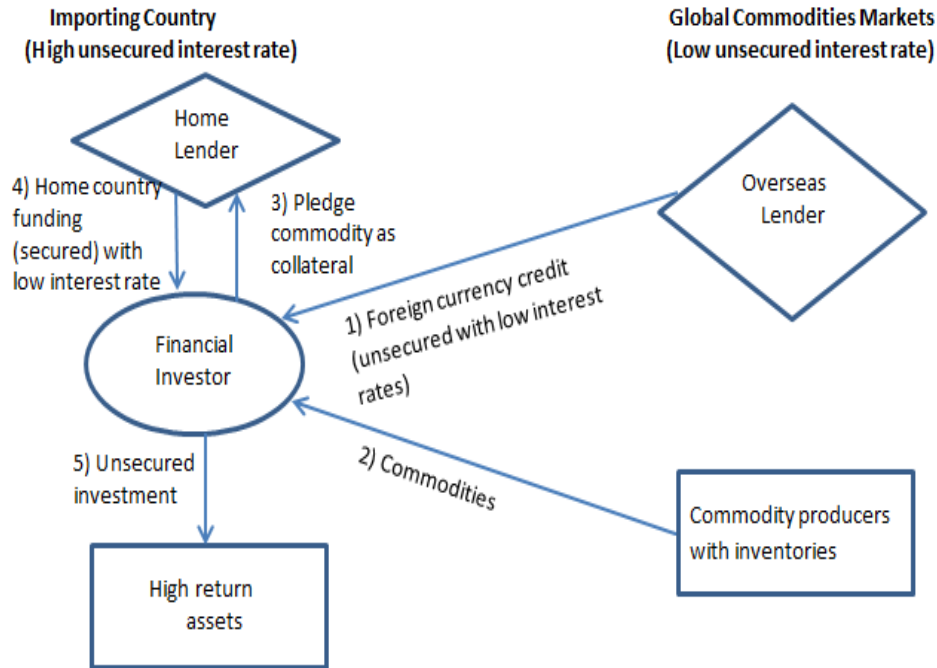
Following [Tang and Zhu \(2016\)](#) carry trade returns are determined by the following main factors: onshore and offshore risk-free interest rate, foreign exchange spot and forward rate. These are the variables we used in this study and are described in the subsequent section. Typically there are two carry trade strategies depending on whether forward contracts are used to implement the trade, they are covered carry trade and uncovered carry trade. When the covered interest rate parity (CIP) holds, the two strategies can be proved to be equivalent. Specifically, in the foreign exchange market, traders set up forward exchange rate according to the CIP, which implies that currencies with a high interest rate are normally traded at a forward discount and currencies with a low interest rate are normally traded at a forward premium ([Cavallo, 2006](#)). Therefore, borrowing

currencies with low interest rates and lending currencies with high interest rates is equivalent to shorting currencies at forward premium and going long currencies at forward discount. In this case, the failure of the UIP indicates that forward rates fail to be unbiased prediction of the future spot rate. Since in practice uncovered carry trade strategy is seldom used, we therefore focus on the covered carry trade return (denoted as  $R_t^c$ ) in this paper, which is calculated as below.

$$R_t^c = \frac{S_t \times (1 + i_t^{on})}{F_t} - i_t^{off} - 1 \quad (1)$$

where  $i_t^{on}$  and  $i_t^{off}$  represent the onshore and offshore risk-free interest rate respectively,  $S_t$  and  $F_t$  stand for the foreign exchange spot and forward rate respectively. The UIP condition indicates that “the expected foreign exchange gain must be just offset by the opportunity cost of holding funds in one currency rather than in the alternative one, measured by the interest rate differential, implying that the expected currency excess returns must be zero” (Laborda et al., 2014, p. 54).

**Fig.1:** Typical process of commodity-based financing



## 4 Data

We consider the classical carry trade pair: USD and CNY. USD has historically been one of the major funding currencies for carry trades due to its low borrowing interest rates and high savings rate as compared to the ones in other developed and some developing countries. Moreover, due to largely expansionary monetary policy and several rounds of quantitative easing, immediately prior and post global financial crisis (GFC) interest rates have been at very low levels in the U.S., which promoted the use of USD to finance investments in China where interest rate was higher.

We use the daily Shanghai 1 month interbank offered rate as a proxy for the onshore risk-free interest rate ( $i_t^{on}$ ), the daily federal funds rate as a proxy for the offshore risk-free interest rate ( $i_t^{off}$ ), the daily CNY to USD exchange rate and CNY to USD 3-month forward rate as a proxy for the foreign exchange spot rate ( $S_t$ ) and forward rate ( $F_t$ ) respectively. The variables of interest

**Table 1:** Summary statistics for variables

Variables	Obs	Mean	Std. Dev.	Min	Max	Skewness	Kurtosis
<b>Panel A: Capital control</b>							
Aluminium price	1307	2001.646	238.687	1525.500	2452.000	-0.262	2.017
Copper price	1307	6519.323	1179.371	4417.500	8724.500	-0.203	1.754
Gold price	2715	38.635	9.974	18.460	61.570	-0.044	2.322
Carry trade return	2715	263.477	223.009	-304.260	951.550	-0.567	3.369
Onshore risk-free rate	2715	355.806	138.235	101.330	969.800	0.574	3.639
Offshore risk-free rate	2715	91.999	164.302	1.000	541.000	1.948	5.144
Foreign exchange spot rate	2715	664.884	46.576	604.120	791.490	0.988	3.319
Foreign exchange forward rate	2715	665.099	43.403	605.500	786.160	0.997	3.432
<b>Panel B: Interest rate control</b>							
Real price of gold	106	284.791	49.806	201.716	402.111	0.553	2.103
Gold spot price	106	267.281	47.376	171.060	371.050	0.266	2.340
Gold futures price (SHFE)	106	267.624	49.273	162.570	384.040	0.240	2.444
Gold futures price (NYMEX)	106	292.139	51.892	177.854	410.979	0.269	2.379
Industrial Production Index	106	110.183	3.703	105.400	121.300	0.693	2.769
Monetary liquidity	106	3.887	1.435	1.659	6.467	0.215	1.793
Gold inventory	106	655000000	783000000	5443110	4380000000	2.320	9.301
Risk premium	106	24.859	5.824	6.794	40.984	0.280	3.352
Foreign exchange spot rate	106	6.488	0.278	6.054	6.950	0.117	1.472
Carry trade return	106	0.035	0.160	0.008	0.080	0.378	2.615

are aluminium, copper and gold price in China and  $R_t^c$  which is calculated using [Eq.\(1\)](#). Data for China's aluminium and copper prices are constructed in three steps: First, we collect the data for aluminium and copper premiums in the Shanghai bonded warehouse (USD per Metric

Tonne). Specifically, they are the premiums paid by customers above the London Metal Exchange (LME) cash aluminium and copper prices. Then, the data for aluminium and copper cash prices (USD per Metric Tonne) in LME is collected. Last, China's aluminium and copper prices are obtained by adding the premiums and the cash prices. Gold price is measured by the China's gold (acceptable purities: 99.99) close price (USD per Gramme) listed in the Shanghai gold exchange (SGE). Apart from the commodity prices, all variables are measured in basis point. Data are obtained from the Thomson Reuters Datastream Database and time span is from October 9th 2006 to March 3rd 2017 with the exception of the aluminium and copper prices (March 1st 2012 to March 3rd 2017). In addition to the data availability issues, this sample period is chosen for two reasons: First, vast amount of research papers have established profitability of carry trade strategies in the environment of high interest rates in target countries. However, after GFC a lot of developed countries which have traditionally been target countries for carry trade investments (e.g., U.S.) initiated expansionary monetary policy in order to stimulate their economies. This has significantly reduced their interest rates, meaning that the spread between the rates in target and funding countries have increased and hence the profitability of carry trade was affected. Following [Aizenman et al. \(2014\)](#), record low interest rates in the U.S. have led to a large scale carry trade activities against high-yielding currencies of emerging economies. Second, although interest rates are still high in the developing countries (e.g., Brazil and China) as compared to their counterparts in the developed countries, they were affected by the GFC as well. In particular, Brazil and China have initiated capital controls to prevent the outflow of capital overseas, which could have also affected the profitability of carry trade strategies.

In addition to performing causality analysis, we estimate [Frankel and Rose \(2010\)](#) model using monthly data. This allows us understanding the impact of carry trade returns on collateral assets prices under the interest rate controls. Due to the data availability, the only commodity we look at is gold. Gold spot price is measured by the China's gold (acceptable purities: 99.99) close price (CNY per Gramme) listed in the SGE. We expressed gold spot price in real term using CPI as a deflator.<sup>3</sup> There are two different measures for gold futures price: one is the continuous trading settlement price (price quotation: Yuan/gramme) on the Shanghai Futures Exchange (SHFE), the other is the continuous trading settlement price (price quotation: USD/per troy ounce) on the New York Mercantile Exchange (NYMEX). For the latter measure, we convert the price quotation into CNY per gramme using the spot exchange rate. Considering the characteristics of gold as a base metal, we choose the industrial production index (IPI) as a proxy for economic activities. To control the impact of rising economic activity on the increase of money supply, we use M2/GDP

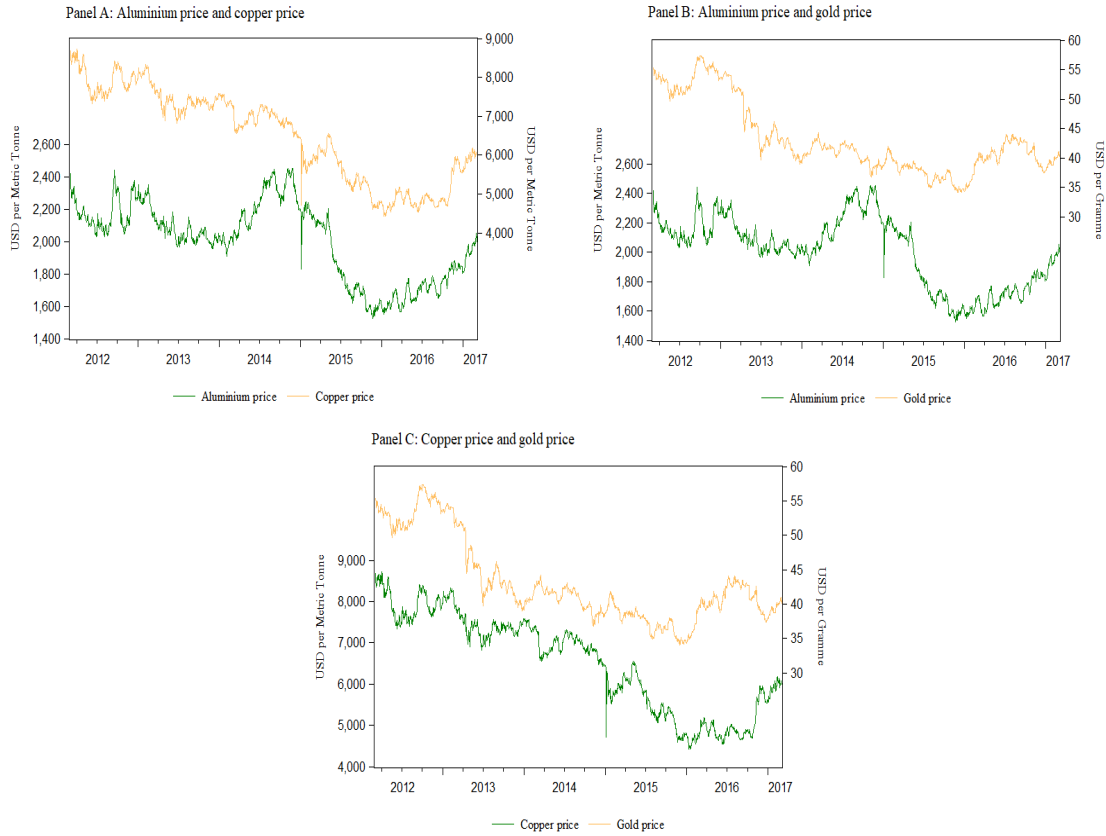
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<sup>3</sup> The formula for adjusting prices for inflation is:  $\frac{P_t}{CPI_t/100}$ , where  $P_t$  is the nominal price of commodity for month  $t$ .

to measure the change in monetary liquidity. Because National Bureau of Statistics of China only provides quarterly GDP data, we convert it into monthly data based on the monthly changing rate of IPI. Gold inventory is measured by the monthly total gold stock<sup>4</sup> of the warehouse (on standard warrant<sup>5</sup>). We use futures-spot spread as a proxy for risk premium. Data on CPI come from Federal Reserve Bank of St. Louis and data for all the other variables are retrieved from Thomson Reuters Datastream Database and time period is from June 2008 to March 2017.

Table 1 presents the definitions and summary statistics for the data organized by the investment country (i.e., China), with additional details including data sources in Appendix.

**Fig.2:** Movement of commodity prices (March 1st, 2012-March 3rd, 2017)



A well-known puzzling phenomenon in financial economics is that the raw commodity prices

<sup>4</sup> We convert the unit from ton to gramme.

<sup>5</sup> In this study, standard warrant refers to the receipt issued by the SHFE certified delivery warehouse to consignor through the SHFE's standard warrant management system for taking delivery of gold.

have a persistent tendency to move together. [Pindyck and Rotemberg \(1990\)](#) found that this price co-movement applies to a broad set of unrelated commodities such as gold, copper, crude oil, lumber, cocoa, wheat and cotton (i.e., cross-price elasticities of demand and supply are close to zero). They considered the possible reason is to some extent due to the herd behaviour in financial markets. [Fig.2](#) plots the movement of prices of the commodities used in this study. Panel A of [Fig.2](#) suggests that the prices of aluminium and copper in China do move together. In contrast, there is no tendency of aluminium and gold prices to react in the same direction (Panel B of [Fig.2](#)). As evident in Panel C, except the mid-year of 2014, no co-movement exist between copper and gold prices. Yet, after the time point, the two commodity prices rise and fall in unison. The observation we made is a reflection of carry trade since both copper and gold are the two most widely used commodities as collateral for carry trade strategies, we argue that this can be a plausible explanation for the price co-movement puzzling. To provide a preliminary glimpse into the relationship between carry trade returns and commodity prices, we plot their behaviours in [Fig.3](#). In general, we observe a clear positive relationship between carry trade returns and all the three commodity prices in China. Moreover, the positive linkage is most significant for the gold price. We consider this is probably because gold is the most frequently used collateral for carry trade.

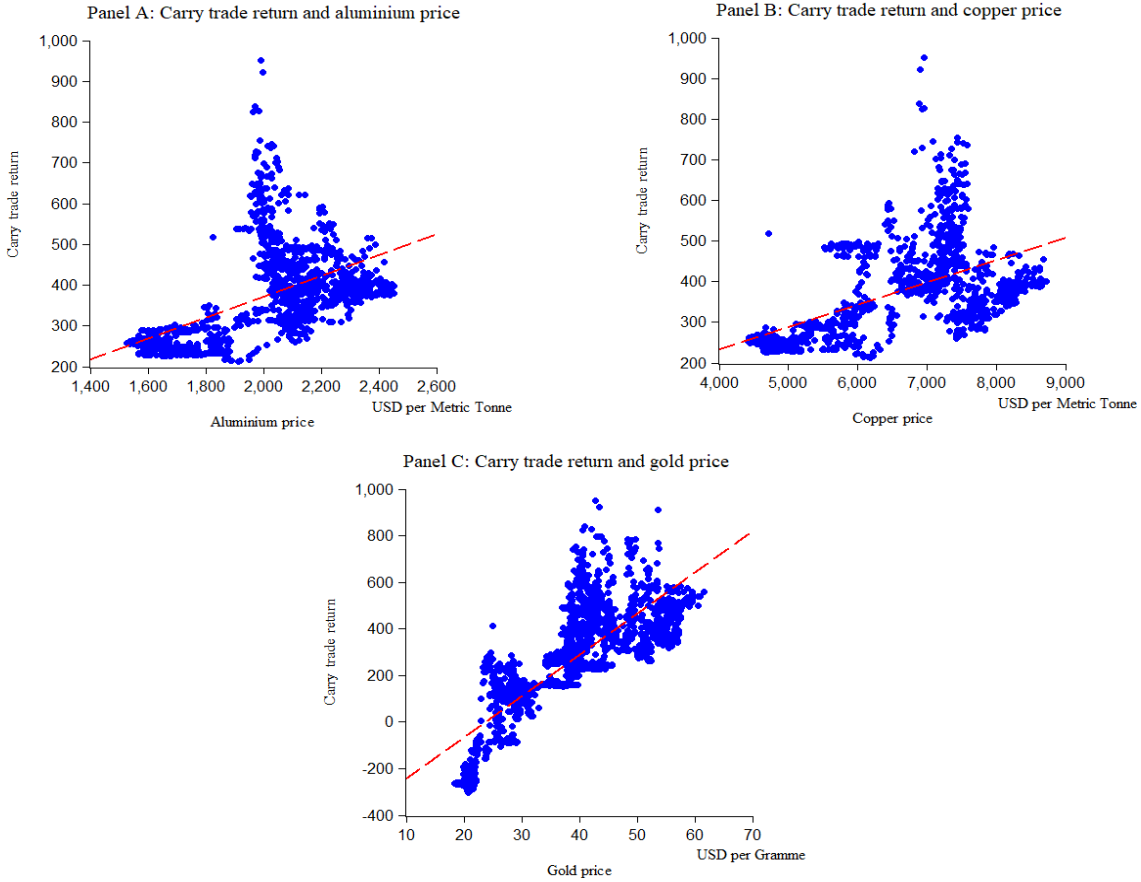
## 5 Empirical Methodology

### 5.1 Tests for unit roots

The natural step towards investigating the long run relationship between returns of carry trade and commodity prices is to first examine the unit root properties of these variables. Moreover, the order of integration and trend specification can affect causality results. The [Hill \(2007\)](#) procedure adopted is also related to these issues. The other issue pertains with the possibility of structural breaks in a VAR model. The presence of structural breaks can cause erroneous results in terms of order of integration in stationarity tests and can further lead to spurious causality results.

It is essential to examine structural breaks in this study. The carry trade returns and commodity prices can be considered to exhibit at least one structural break during the sample period. Specifically, the 2007-2008 GFC and the 2015 Chinese stock market crash. In this paper, instead of assuming exogenous structural breaks, we apply methods in which the breakpoints are estimated rather than fixed.

**Fig.3:** Carry trade returns and commodity prices



*Note:* Time span of Panel A and B is from March 1st 2012 to March 3rd 2017, while Panel C is from October 9th 2006 to March 3rd 2017.

The present study is built in two directions based on these aspects. First, we base our analysis on the traditional unit root tests, which are commonly used in the literature, and further investigate the trending nature of carry trade returns and commodity prices. Then, we consider the possibility of structural breaks when testing for unit roots using the latest methods of [Narayan et al. \(2016\)](#). The test allows us to examine the presence of structural breaks and stationarity properties of the time series data when the noise component can be either stationary or integrated.

We first employ a number of unit root tests that are frequently used in the relevant studies to test for the null of non-stationarity against trend stationary alternatives. In particular, we apply the augmented [Dickey and Fuller \(1979\)](#), ADF test, the [Phillips and Perron \(1988\)](#), PP test), the KPSS ([Kwiatkowski et al., 1992](#)) test, the GLS transformed Dickey-Fuller ([Elliott et al., 1996](#),



(1990) test, the ARDL model has more power. Second,  $I(0)$  variables are allowed in the ARDL model. Third, it is easy to interpret the ARDL model since it has only one single equation. Fourth, as argued in Laurenceson and Chiai (2003), the ARDL model uses a sufficient number of lags to capture the data-generating process in a general-to-specific modeling framework. Last, the ARDL model is able to manage both long-run cointegration and short-run dynamics.

The specification of our ARDL model with structural breaks is as follows:

$$CR_t = \alpha_0 + \sum_{i=1}^j \beta_j AP_{t-j} + \sum_{i=1}^k \gamma_k CP_{t-k} + \sum_{i=1}^l \delta_l GP_{t-l} + \lambda_i B_{it} + \varepsilon_t \quad (4)$$

where  $CR$ ,  $AP$ ,  $CP$  and  $GP$  stand for carry trade returns, aluminium price, copper price and gold price respectively,  $B_{it}$  ( $i = 1, 2$ ) is the break dummy which equals one at the two break points that are identified by the Narayan et al. (2016) unit root test for all variables. The terms  $j$ ,  $k$  and  $l$  are number of lags of the independent variables. The optimal number of lags are decided by the information criterion. The ARDL approach estimates  $(p + 1)^q$  equations<sup>7</sup> to obtain the optimal lags for each variable, where  $p$  denotes the maximum number of lags and  $q$  represents the number of regressors.

After deciding the number of lags in the model, we establish the unrestricted error correction model (ECM) as below:

$$\begin{aligned} \Delta CR_t = \alpha + \sum_{i=1}^j \beta_j \Delta CR_{t-j} + \sum_{i=1}^k \gamma_k \Delta AP_{t-k} + \sum_{i=1}^l \delta_l \Delta CP_{t-l} + \sum_{i=1}^m \Delta GP_{t-m} \\ + \lambda_i B_{it} + \mu_0 CR_{t-1} + \mu_1 AP_{t-1} + \mu_2 CP_{t-1} + \mu_3 GP_{t-1} + \varepsilon_t \end{aligned} \quad (5)$$

Next, a restricted ECM is used to investigate the short-run dynamics based on the results of bounds test. We conduct the analysis by following the steps below: First, we lag the residuals from Eq.(4) by one period. Then the lagged residual is added to Eq.(5) as the error correction term to construct the restricted ECM. The ARDL<sup>8</sup>(5, 0, 0, 0, 0) restricted ECM can be specified as follows:

$$\begin{aligned} \Delta CR_t = \alpha + \sum_{i=1}^j \beta_j \Delta CR_{t-j} + \sum_{i=1}^k \gamma_k \Delta AP_{t-k} + \sum_{i=1}^l \delta_l \Delta CP_{t-l} + \sum_{i=1}^m \Delta GP_{t-m} \\ + \lambda_i B_{it} + \varphi ECT_{t-1} + \varepsilon_t \end{aligned} \quad (6)$$

where ECT denotes the error correction term.

<sup>7</sup> In this study, we estimate 52488 equations in total.

<sup>8</sup> The decision procedure is presented in Section 6.

### 5.3 Testing for long-run non-causality in VAR processes

Given the structural breaks and crises identified in the financial time series, non-linear causal relationship is likely to exist due to volatility and return spillovers. Because the linear and non-linear causal relationships are dependent to the sample data, we adopt a causality framework with dynamic rolling window. Specifically, the Hill (2007) fixed-length rolling window causality test is used. In this section, we briefly introduce the test. The procedure is based on Wald type test statistic with the joint null hypotheses of zero parameter linear restrictions. They are developed in a VAR system of order  $p$  as below:

$$W_t = \mu_t + \sum_{k=1}^p \pi_k W_{t-k} + a_t, \quad t = 1, \dots, T \quad (7)$$

where  $W_t = (w_{1,t}, w_{2,t}, \dots, w_{m,t})'$  is an  $m \times 1$  random vector with possibly integrated series of order at most  $d$ ,  $\mu_t$  denotes a deterministic trend and its most common form includes only the constant term,  $\mu_t = \mu$ , although trends, seasonal or other type of dummies can also be considered;  $\pi_k$  represents  $m \times m$  coefficient matrices and  $a_t$  is an  $m \times 1$  vector white noise process with nonsingular covariance matrix  $\Omega = E(a_t a_t')$ .

#### 5.3.1 Hill (2007) efficient test of long-run causality

Based on the Dufour and Renault (1998) framework and the generalization of the standard definition of Granger causality, Hill (2007) proposed a recursive parametric representation test procedure for examine multi-step ahead causation in trivariate VAR system (i.e.,  $X$ ,  $Y$  and an auxiliary variable  $Z$ ), which can be applied to capture causality chains. Causal chains can let multi-period causation delays, that is, periods of non-causation can be followed by causation. Hence, the test helps in understanding the impact of one variable on another as ‘‘Direct’’ or ‘‘Indirect through an auxiliary variable’’ (i.e.,  $Y$  causes  $Z$  and further  $Z$  causes  $X$ ). Moreover, this feature enables to provide useful insights when investigate causal relations, given the sluggishness of macroeconomic time series variables.

The testing procedure is based on the estimation of a VAR model as in Eq.(7) and use nonlinear recursive representations for the coefficients  $\pi_k^{(h)}$  in the Dufour et al. (2006) framework<sup>9</sup> that

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<sup>9</sup>The point of origin is the ordinary least square (OLS) estimation of the following autoregression of order  $p$  at horizon  $h$ , which is named  $(p, h)$ -autoregression by Dufour et al. (2006):

$$W_{t+h} = \mu + \sum_{k=1}^p \pi_k^{(h)} W_{t+1-k} + u_{t+h}^{(h)}$$

In the Dufour et al. (2006) approach, the VAR( $p$ ) process in Eq.(7) is an autoregression at horizon one, and the above equation is a projection of Eq.(7) at any horizon  $h$ , given the available information at time  $t$ . Dufour and Renault (1998) provided formulas of the coefficients  $\pi_k^{(h)}$  (see Eq.(3.7), (3.8), (3.16) and (3.17)) and also the  $(p, h)$ -autoregression in matrix form.

“lead” back to the VAR model parameters  $\pi_k$  (see [Dufour and Renault, 1998](#), lemma 3.1, p.1109 and [Hill, 2007](#), lemma 3.1, p.752). [Hill \(2007\)](#) introduced a recursive parametric representation of causality chains for the trivariate VAR processes case.

[Hill \(2007\)](#) showed that a causality chain from variable  $Y$  to  $X$  through  $Z$  indicates that  $Y$  will eventually cause  $X$  if  $Z$  is univariate, given the linear necessary and sufficient conditions for non-causation up to arbitrary horizons (see Theorem 2.1.iv). The procedure also considers the possibility of cointegration in the VAR system, since the [Toda and Yamamoto \(1995\)](#) and [Dolado and Lütkepohl \(1996\)](#) augmented lags approach is applicable.

The sequential testing procedure<sup>10</sup> has three steps with Wald type tests performed in each of them. First, we examine if variable  $Y$  never causes variable  $X$  and  $Z$  (i.e., one-step or multi-steps ahead), and similarly  $Y$  and  $Z$  does not cause  $X$ . The rejection of both hypotheses (*test* 0.1, 0.2) indicates to test for horizon-specific non-causality. The notation from [Hill \(2007\)](#) of the hypothesis testing is as below:

$$H_0^{(\infty)} : Y \not\rightarrow (X, Z) \iff \pi_{XY} = \pi_{ZY} = 0 \quad (\text{test 0.1})$$

$$H_0^{(\infty)} : (Y, Z) \not\rightarrow X \iff \pi_{XY} = \pi_{XZ} = 0 \quad (\text{test 0.2})$$

The second step can be divided into two stages. First, testing if  $Y$  does not cause  $X$  one-step ahead (*test* 1.0). If there is no direct causal relationship between  $Y$  and  $X$ , we then perform intermediary tests to examine the existence of a causal chain through  $Z$  (*tests* 1.1, 1.2). If either hypothesis cannot be rejected, a broken causal chain is obtained and it can be concluded that  $Y$  never causes  $X$  at any horizon  $h > 0$ . Using the hypothesis testing notation from [Hill \(2007\)](#), we have:

$$H_0^{(1)} : Y \not\rightarrow X \iff \pi_{XY} = 0 \quad (\text{test 1.0})$$

$$H_0^{(1.1)} : Y \not\rightarrow Z \iff \pi_{ZY} = 0 \quad (\text{test 1.1})$$

$$H_0^{(1.2)} : Z \not\rightarrow X \iff \pi_{XZ} = 0 \quad (\text{test 1.2})$$

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<sup>10</sup> Following [Hill \(2007\)](#), we define “ $Y$  does not cause  $X$  at horizon  $h > 0$ ” (denoted  $Y \not\rightarrow^h X|I_{XZ}$ , where  $I_{XZ}$  represents the set of information common to all periods and contained in the past and present  $X$  and  $Z$ .) if incorporating past and present values of  $Y$  does not improve the minimum mean-squared-error forecast of  $X_{t+h}$  for any  $t$ . We say “ $Y$  does not cause  $X$  up to horizon  $h > 0$ ” (denoted  $Y \not\rightarrow^{(h)} X|I_{XZ}$ ) if  $Y \not\rightarrow^h X|I_{XZ}$  for each  $k = 1, \dots, h$ . Finally, we define “ $Y$  does not cause  $X$  at any horizon  $h > 0$ ” (denoted  $Y \not\rightarrow^\infty X|I_{XZ}$ ) if  $Y \not\rightarrow^h X|I_{XZ}$  for every  $h > 0$ .

If both are rejected then we proceed to the third step.<sup>11</sup> For the third also the last step, if a causal chain is found, then non-causality up to horizon  $h \geq 2$  is tested. Using the notation from Hill (2007) of the hypothesis testing, the testing sequence has the following form:

$$H_0^{(h)} : Y \not\rightarrow X \iff \pi_{XY} = \pi_{XZ,i} = 0, \quad i = 1, \dots, h-1 \quad (\text{test } h.0)$$

Under weak regularity conditions, the Wald-type statistics are used to test all hypotheses discussed above that converge asymptotically to  $\chi^2$  variates. However, similar to the Dufour et al. (2006) Wald tests, the  $\chi^2$  distribution can be a poor proxy for the true small sample distributions. Hill (2007) developed a parametric bootstrap approach for simulating small sample  $p$ -values and is applied in this paper. The Hill's approach needs Bonferroni-type test size bounds to control the overall size of the tests and the procedure is discussed in detail in Hill (2007, p. 756).

## 6 Results and discussion of findings

Table 2 presents the results of the standard unit root tests for all variables of interest. The evidence in favour of non-stationarity in levels is overwhelming. However, the results for the first differenced variables suggest that the variables are stationary at the 5% significance level or better. Overall, we conclude that all series are integrated of order one,  $I(1)$ .

The results of Narayan et al. (2016) unit root test are reported in Table 3. There is evidence of mean reversion in copper price and carry trade returns. In contrast, for aluminium and gold prices, the null hypothesis of unit root cannot be rejected at the 5% significance level or better. In terms of the estimated breaks, we notice both two breaks in China's commodity market appears in the post GFC period (first break: early 2013 and early 2015; second break: mid 2013 and early 2015). Nevertheless, the first (Feb-2008) and second (Jun-2008) breaks are detected during the GFC period in the carry trade return series. All the identified breaks can be linked with the major domestic or international shocks that affected China's commodity market. As evident, the breaks occurred between February 2008 and June 2008 is related to the GFC, suggesting that it had a significant influence on Chinese carry traders. Specifically, after July 2008, the credit crunch induced a sudden and unexpected unwinding of the dollar carry trade which is very important for China, leading to a sharp appreciation in the dollar, carried the Yuan which is pegged to the dollar, upward with it. Moreover, both U.S. and China are uncomfortably poised between

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<sup>11</sup> Notice that this step is reached only if evidence suggests non-causation  $Y \not\rightarrow X$  and a causal chain  $Y \rightarrow Z \rightarrow X$ .

**Table 2:** Standard unit root tests results

Variables	Tests								
	ADF	PP	KPSS	DF-GLS	$MZ_{\alpha}^{GLS}$	$MZ_t^{GLS}$	MSB	MPT	ERS P.O
<b>Panel A: capital control</b>									
<b>Levels</b>									
Aluminium price	-2.010	-2.127	2.355***	-0.577	-0.934	-0.576	0.616	20.641	21.299
Copper price	-1.690	-1.925	3.876***	0.052	0.054	0.052	0.958	52.631	54.522
Gold price	-2.032	-2.002	2.982***	-0.061	-0.062	-0.061	0.983	53.514	54.532
Carry trade return	-2.638*	-2.840*	3.406***	-0.341	-0.508	-0.348	0.686	26.686	25.980
<b>First Differences</b>									
d(Aluminium price)	-10.810***	-41.402***	0.121	-1.379	-2.673	-1.152	0.431	9.151	0.205***
d(Copper price)	-9.632***	-46.978***	0.126	-2.565**	-2.805	-1.174	0.419	8.705	0.282***
d(Gold price)	-16.378***	-54.710***	0.317	-10.136***	-192.543***	-9.803***	0.051***	0.140***	0.044***
d(Carry trade return)	-27.687***	-45.361***	0.104	-7.434***	-24.342***	-3.487***	0.143***	1.011***	0.020***
<b>Panel B: interest rate control</b>									
<b>Levels</b>									
gold real price	-1.629	-1.678	0.217***	-1.211	-2.815	-1.178	0.419	32.125	34.697
monetary liquidity	-2.068	-2.083	0.289***	-1.266	-3.963	-1.319	0.333	21.945	116.123
industrial production index	-3.229*	-3.343*	0.137*	-2.548	-12.591	-2.488	0.198	7.357	5.511**
gold inventory	-1.884	-6.283***	0.120*	-1.842	-4.060	-1.378	0.339	21.922	19.872
risk premium	-2.768	-6.191***	0.226***	-2.460	-9.835	-2.204	0.224	9.328	18.903
one period lagged futures price	-1.643	-1.583	0.243***	-1.318	-3.363	-1.274	0.379	26.657	28.501
carry trade return	-2.024	-2.815*	0.329	-1.099	-2.920	-1.171*	0.401	8.304	8.374
<b>First Differences</b>									
d(gold real price)	-2.162	-12.307***	0.133*	-1.635	-0.892	-0.644	0.722	95.990	31.074

*Continued on next page*

Table 2 – *Continued from previous page*

d(monetary liquidity)	-11.016***	-13.831***	0.103	-10.968***	-51.116***	-5.036***	0.099***	1.881***	1.876***
d(industrial production index)	-9.101***	-10.212***	0.055	-1.818	-1.354	-0.807	0.596	65.281	2.135***
d(gold inventory)	-13.564***	-27.905***	0.321***	-13.653***	-47.087***	-4.852***	0.103***	1.936***	1.910***
d(risk premium)	-17.393***	-33.453***	0.119*	-17.507***	-38.582***	-4.391***	0.114***	2.369***	2.336***
d(one period lagged futures price)	-3.131	-11.307***	0.113	-3.150**	-10.748	-2.309	0.215	8.523	8.076
d(carry trade return)	-11.758***	-16.977***	0.269	-11.718***	-50.378***	-5.019***	0.100***	0.486***	0.483***

*Note:* The KPSS test has the null hypothesis of stationarity. For all other tests, the null hypothesis is there is a unit root in the series. As suggested by [Ng and Perron \(2001\)](#), the lag length for the ADF, DF-GLS,  $MZ_{\alpha}^{GLS}$ ,  $MZ_t^{GLS}$ , MSB, MPT and ERS P.O tests is selected using the modified Akaike information criterion (MAIC). The PP and KPSS tests use the automatic bandwidth selection technique of Newey-West using Bartlett Kernel computing the spectrum.

\*, \*\*, \*\*\* Denote statistical significance at the 10%, 5% and 1% level respectively.

**Table 3:** Narayan et al. (2016) unit root test with two structural breaks

Variables	Test Statistic	TB1	TB2
Aluminum price	-2.563	Jan-15, 2015	Feb-09, 2015
Copper price	-4.198*	Jan-07, 2013	Jan-05, 2015
Gold price	-3.316	Apr-01, 2013	May-03, 2013
Carry trade return	-30.795*	Feb-15, 2008	Jun-10, 2008

*Note:* TB1 and TB2 denotes dates of structural breaks. The 5% critical value for the unit root test statistic is -3.76, obtained from Narayan et al. (2016) [Table 3 for  $N = 250$  and GARCH parameters  $[\alpha, \beta]$  chosen as  $[0.05, 0.90]$ ]. Narayan et al. (2016) only provided critical values for 5% significance level.

\* Denotes statistical significance at the 5% level.

inflation and deflation. The GFC in 2008 attacked hard and forced carry traders in dollar, yen and commodities to unwind their positions. Prior to this date, the major threat comes from inflation due to the volatility of international commodity prices and the internal loss of monetary control from the one-way bet that the value of Yuan always appreciates. Although being partly endogenous, the accidental decline in commodity prices was also a partly exogenous deflationary shock to the global economy. At the end of 2008, the crisis became much greater and was characterized by the sudden collapse in commodity prices. The volume of international trade fell substantially. In particular, China's exports dropped by half from mid-2008 into 2009. The break for the commodity prices series in the early 2015 can be associated with China's economic slowdown. China has been considered as a global consumer of commodities. Commodity prices tend to rise when the economy booms and fall when it falters. There has been a clear correlation between Chinese GDP growth and commodity prices. The annual growth for China in 2014 was the slowest since 1990. Accompanying the slowdown is China's cooling demand for commodities, both domestically and internationally. Many commodity prices started to drop since 2012.<sup>12</sup> Several economists recognized that a global output surplus in many commodity markets and a further declining in demand from China contributed to the underpin market weakness and drive prices down. The stationary property of China's carry trade market after allowing structural breaks reflects the timely intervention of Chinese government in responds to the GFC. Specifically, to prevent further appreciation of Yuan, the People's Bank of China (PBC) in early July 2008 reset the Yuan/dollar rate at  $6.83 + 0.3$  percent which remained for almost one year. The refixed

<sup>12</sup> For example, energy prices have fallen by 70% and metals prices by 50%.

Yuan/Dollar rate had a dramatic influence on China’s financial markets. Net hot money inflows stopped due to the one-way bet on exchange appreciation had ended. Furthermore, private financial capital began to flow outward to finance China’s huge current account surplus of more than \$300 billion each year. Also after the PBC getting their internal monetary controlled, owing to the sharp decrease in exports they focused on domestic credit expansion. In particular, they cut domestic reserve requirements on commercial banks and loosened other direct constraint on bank lending. In addition, the lending rates remained about 3 percentage points higher than the deposit rates to keep banks’ profitability.

We use four methods to decide the optimal ARDL model: the Akaike Information Criterion (AIC), Schwartz Bayesian Criterion (SBC), Hannan-Quinn Criterion (HQ) and adjusted R-squared. Considering almost all factors (e.g., significance of coefficients, goodness of fit of the model, serial correlation, stability of the model), HQ method is adopted to select the ARDL(5, 0, 0, 0, 0) as our benchmark specification. We then apply the Breusch-Godfrey Serial Correlation Lagrange Multiplier (LM) test to examine whether the ARDL(5, 0, 0, 0, 0) model is free of serial correlation. Panel A of Table 4 reports the LM test results. Both the F-statistic and observed R-squared statistic cannot reject the null hypothesis of no serial correlation<sup>13</sup> in the unrestricted ECM at the 5% significance level, indicating that there is no serial correlation in the residual. The Cusum test is used to verify the model stability and the result shows that our model is stable at the 5% significance level (Panel A of Fig.A1 in Appendix). After confirming our model has neither serial correlation nor instability, we proceed to the bounds test.

**Table 4:** ARDL unrestricted error correction model with structural breaks

<b>Panel A:</b> Breusch-Godfrey Serial Correlation LM test				
Test Statistic		<i>p</i> -values		
F-statistic	0.151	0.860		
Observed R-squared	0.305	0.859		
<b>Panel B:</b> Bounds test				
Test Statistic		Lower Bound	Upper Bound	
F-statistic	5.774	2.56	3.49	
<b>Panel C:</b> Long-run coefficients				
Variables	Coefficient	Std. Error	<i>t</i> -statistic	<i>p</i> -values
Aluminium price	0.062	0.099	0.626	0.531
Copper price	0.090	0.026	3.495	0.001
Gold price	-11.434	3.395	-3.368	0.001
Break dummy	88.000	171.219	0.456	0.649

*Note:* The lower bound and upper bound listed in the table are the 5% significance level critical value bounds.

<sup>13</sup> Our results are robust from 1 lag to 10 lags. The result reported in Table 4 is the case of 2 lags.

The Pesaran et al. (2001) bounds test is applied to examine the long-run equilibrium between returns of carry trade and commodity prices. Specifically, it is an F-test which has the null hypothesis that  $\mu_0 = \mu_1 = \mu_2 = \mu_3 = 0$  in Eq.(5). According to Pesaran et al. (2001), the lower bound is used when all variables are  $I(0)$ , and the upper bound is used when all variables are  $I(1)$ . There is likely to have no cointegration between carry trade returns and commodity prices if the F-statistic is below the lower bound. If the F-statistic is higher than the upper bound, indicating the existence of cointegration. Alternatively, the evidence of cointegration is ambiguous when the F-statistic falls between the lower bound and upper bound. Panel B of Table 4 presents results of the bounds test. It can be seen that the F-statistic is larger than the upper bound at the 5% significance level, suggesting that there is a long-run association between carry trade returns and commodity prices.

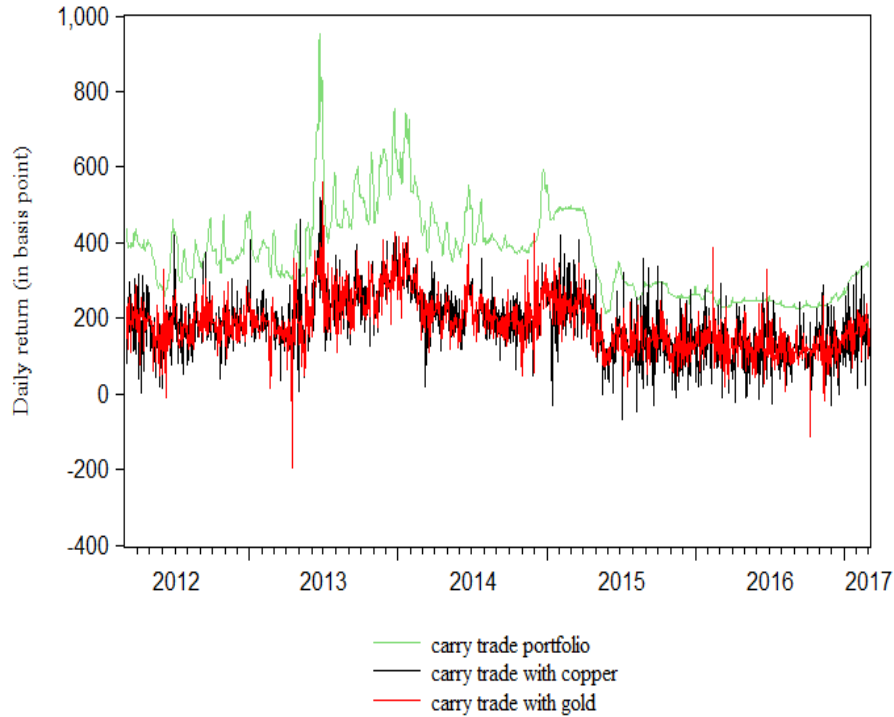
Next we extract a long-run multiplier between the dependent and independent variables from the unrestricted ECM to obtain the long-run coefficients in Eq.(4). As evident, Panel C of Table 4 shows that the coefficient of copper price is significant and assumes a positive value this indicates a positive relationship between the copper price and carry trade returns. Specifically, in the long-run, 1 dollar increase in copper price can lead to a 0.09 basis point rise in carry trade returns. In contrast, there is an inverse relationship between gold price and returns of carry trade. In particular, 1 dollar increase in gold price would yield a decrease of 11.43 basis point in returns of carry trade. We do not find a statistically significant relationship between aluminium price and carry trade returns. Our findings are possibly owing to the fact that copper and gold are widely used as collateral in China for carry trade. Moreover, the small positive sign of copper price and the large negative sign of gold price indicate that in long-run there are hedge characteristics for copper and gold returns on carry trade returns. Table 5 compares three different portfolios includes a full carry trade portfolio, one portfolio consists of 50% carry trade and 50% copper and another portfolio contains half carry trade and half gold. The comparison shows that adding copper and gold in the carry trade portfolio reduces the standard deviation. Furthermore, compared with the copper portfolio, the Sharpe ratio over the time period including gold in the portfolio provides a higher payoff given the risk undertaken. Fig.4 illustrates that the full carry trade portfolio outperforms the copper and gold portfolios over time, yet this comes at a cost of high risk. In addition, the copper and gold portfolios have a similar return over the whole sample period, but the volatility is substantially larger for the former after mid 2014.

**Table 5:** Comparison of trading strategies (Mar 2nd, 2012 to Mar 3rd 2017)

Portfolio	Average Daily Return	Std. Dev	Sharpe Ratio	% Carry Trade	% Gold
Carry trade portfolio	371.325	115.839	3.042	100%	0%
Carry trade portfolio with copper	184.932	101.924	1.628	50%	50%
Carry trade portfolio with gold	184.645	72.906	2.272	50%	50%

*Note:* All returns are authors' calculation and measured in basis point. We also compute the average daily federal funds rate as the risk free rate to obtain the Sharpe ratio.

**Fig.4:** Portfolio simulation (Mar 2nd, 2012 to Mar 3rd, 2017)



Then the restricted ECM is used to estimate the short-run coefficients. ARDL(8, 0, 0, 4, 0) is selected based on the AIC criterion. We again need to check whether the model passes the serial correlation test and model stability test before estimating results. Panel A of Table 6 presents the serial correlation test results of the ARDL restricted ECM. Both two test statistics cannot reject the null hypothesis, suggesting that our model has no serial correlation. Moreover, the

stability of the restricted ECM is confirmed by the Cusum test (Panel B of Fig.A1 in Appendix). The estimation results of the restricted ECM is outlined in Panel B of [Table 6](#). The short-run coefficients for aluminum and copper prices are not significant, which implies that there are no short-term relationships between the prices of the two commodities and carry trade returns. We observe a mix of positive and negative signs in the coefficients of gold price. Specifically, only the second and third lag are statistically significant at the 5% significance level for the coefficients of gold price. The coefficient for second lag is the largest in scale, at 3.20, meaning that 1 dollar rise in gold price can increase the returns of carry trade around 3.20 basis point, and it can take two

**Table 6:** ARDL restricted error correction model with structural breaks

<b>Panel A:</b> Breusch-Godfrey Serial Correlation LM test				
Test Statistic	<i>p</i> -values			
F-statistic	0.638	0.528		
Observed R-squared	1.295	0.523		
<b>Panel B:</b> Short-run coefficients				
Variables	Coefficient	Std. Error	<i>t</i> -statistic	<i>p</i> -values
C	-0.065	0.507	-0.129	0.898
Break dummy	2.737	6.587	0.416	0.678
D(Carry trade return(-1))	0.202	0.028	7.228	0.000
D(Carry trade return(-2))	-0.020	0.028	-0.694	0.488
D(Carry trade return(-3))	0.064	0.028	2.271	0.023
D(Carry trade return(-4))	0.087	0.028	3.086	0.002
D(Carry trade return(-5))	-0.015	0.028	-0.535	0.593
D(Carry trade return(-6))	0.014	0.028	0.480	0.631
D(Carry trade return(-7))	-0.074	0.028	-2.606	0.009
D(Carry trade return(-8))	-0.061	0.028	-2.189	0.029
D(Aluminium price)	-0.005	0.032	-0.154	0.877
D(Copper price)	0.003	0.008	0.401	0.688
D(Gold price)	-0.534	1.281	-0.417	0.677
D(Gold price(-1))	-0.803	1.268	-0.634	0.526
D(Gold price(-2))	3.195	1.267	2.522	0.012
D(Gold price(-3))	-2.163	1.266	-1.708	0.088
D(Gold price(-4))	-1.982	1.267	-1.565	0.118
ECT(-1)	-0.032	0.007	-4.553	0.000

*Note:* C represents the constant term. D before each variable stands for the first difference operator and the numbers in the parenthesis behind each variable are the number of lags taken. ECT denotes the error correction term.

days to have such influence. Yet, the impact become negative on the third day which leads to 2.16 basis point declining in carry trade returns. The last row in [Table 6](#) reports the coefficient of

error correction term which is between 0 and -1 and is statistically negative, ensuring convergence to a significant long-run relationship. The coefficient -3.20% refers to the speed of adjustment to the long-run equilibrium, which implies that nearly 3% of any disequilibrium from the long-run is corrected within one period, that is, one day for our data.

We do not need the cointegration specification because using the augmented lags method suggested by [Toda and Yamamoto \(1995\)](#) and [Dolado and Lütkepohl \(1996\)](#), the [Hill \(2007\)](#) approach for testing non-causality can be directly applied on a VAR( $p$ ) process in levels. Following this way, we augment the lag order of VAR model by  $d$  extra lags, where  $d$  denotes the maximum order of integration, Wald type restrictions can be imposed only on the first  $p$  coefficient matrices and the test statistics follow standard asymptotic distributions. [Dufour et al. \(2006\)](#) showed that this extension can be applied to examine non-causality at different time horizons based on standard asymptotic theory in non-stationary and cointegrated VAR systems without pre-specifying the cointegration relationships. As far as the maximal order of integration does not exceed the true lag length of the model, the conventional lag selection procedure then can be employed to a possibly integrated or cointegrated VAR model. [Table 7](#) tabulates the results of the lag order selection for all variables. Following the rule of [Dufour and Renault \(1998\)](#), non-causality is tested up to horizon  $h = p + 1$  where  $p$  denotes the number of lags in VAR model.

**Table 7:** Lag order selection for VAR model

Criteria					Selection
LR	FPE	AIC	SC	HQ	
8	2	2	2	2	2

*Note:* LR: sequential modified likelihood ratio statistic; FPE: Final Prediction Error; AIC: Akaike Information Criterion; SC: Schwarz Criterion; HQ: Hannan & Quinn Criterion.

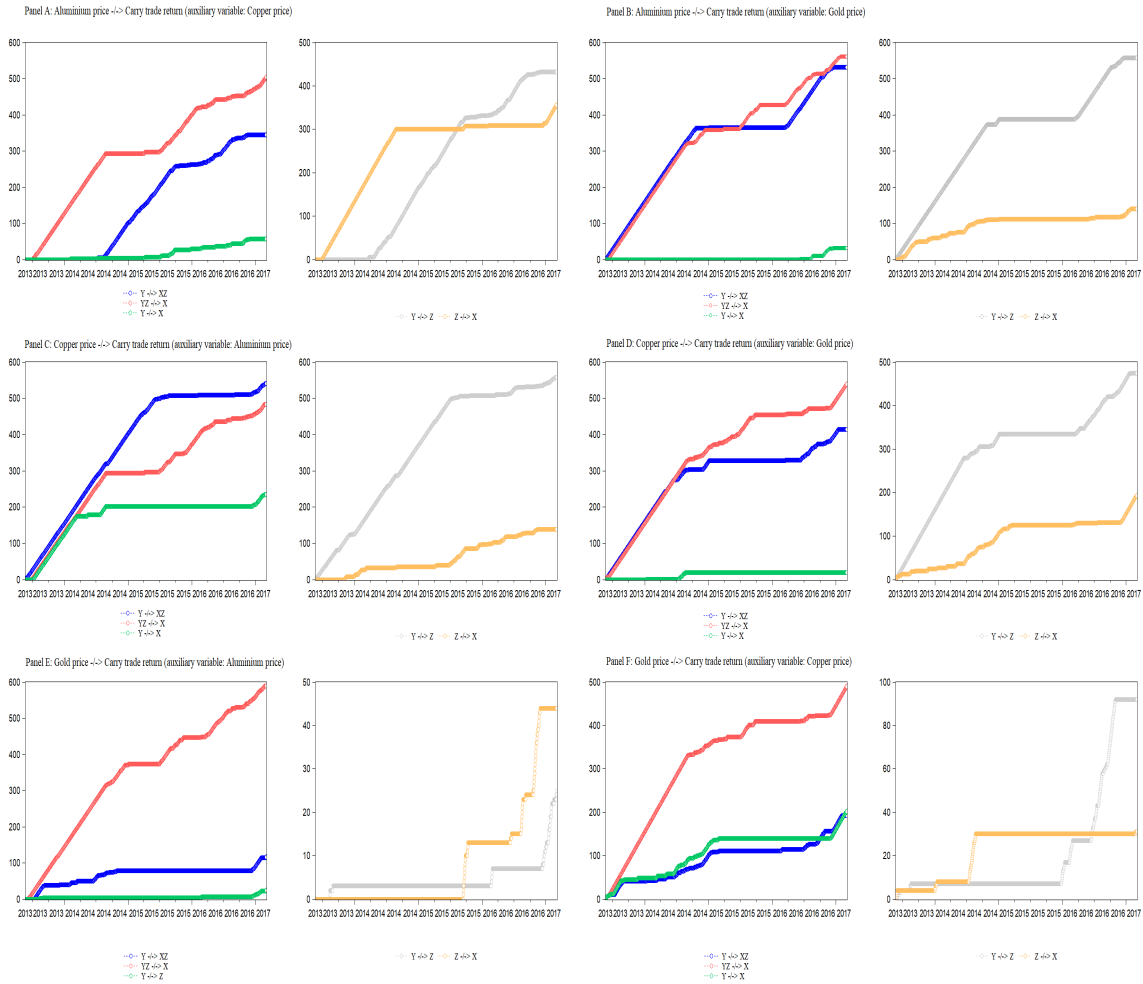
The results of [Hill \(2007\)](#) test are reported in [Table 8](#). The criterion for detection of non-causality at all horizons ( $Y \not\rightarrow^{\infty} X$ ) is a failure to reject either test 0.1 or test 0.2. We reject at horizon one if we reject  $Y \not\rightarrow^1 X$ ; we reject  $Y \not\rightarrow^2 X$  if we fail to reject  $Y \not\rightarrow^1 X$ , reject both intermediary tests (test 1.1 and test 1.2), and reject test 2.0; and so on. We do not allow for rejection at multiple horizons for a particular window. If we reject  $Y \not\rightarrow^h X$ , we stop the test procedure for the specific window. In this sense, we concern the earliest horizon at which causation appears. We do, however, allow for detection of non-causality at all horizons (i.e.,  $Y \not\rightarrow^{\infty} X$ ) and causality at some horizons (i.e.,  $Y \xrightarrow{h} X$ ). We present window frequencies in which the two sets

**Table 8:** Hill (2007) test results

Causality Direction	Auxiliary Variable	Avg. VAR order	Avg. $p$ -values	Tests						
				0.1	0.2	1.0	1.1	1.2	2.0	3.0
<b>Panel A:</b> Testing from commodity prices to carry trade returns										
Aluminium price $\nrightarrow$ Carry trade return	Copper price	2.961	0.308	0.566	0.519	0.695	0.299	0.259		
Aluminium price $\nrightarrow$ Carry trade return	Gold price	2.685	0.206	<b>0.000</b>	0.266	0.840	<b>0.000</b>	0.100		
Copper price $\nrightarrow$ Carry trade return	Aluminium price	2.961	0.308	<b>0.000</b>	0.480	0.359	<b>0.000</b>	0.530		
Copper price $\nrightarrow$ Carry trade return	Gold price	2.940	0.172	<b>0.000</b>	0.426	0.417	<b>0.000</b>	0.419		
Gold price $\nrightarrow$ Carry trade return	Aluminium price	2.685	0.206	0.522	0.790	0.393	0.592	0.624		
Gold price $\nrightarrow$ Carry trade return	Copper price	2.940	0.172	0.445	0.403	0.248	0.586	0.658		
<b>Panel B:</b> Testing from carry trade returns to commodity prices										
Carry trade return $\nrightarrow$ Aluminium price	Copper price	2.961	0.308	0.319	<b>0.000</b>	0.210	<b>0.100</b>	<b>0.000</b>		
Carry trade return $\nrightarrow$ Aluminium price	Gold price	2.685	0.206	0.907	0.573	0.645	0.962	0.356		
Carry trade return $\nrightarrow$ Copper price	Aluminium price	2.961	0.308	0.407	<b>0.000</b>	0.142	0.263	0.341		
Carry trade return $\nrightarrow$ Copper price	Gold price	2.940	0.172	0.715	<b>0.000</b>	0.428	0.918	0.792		
Carry trade return $\nrightarrow$ Gold price	Aluminium price	2.685	0.206	0.922	<b>0.000</b>	0.972	0.664	<b>0.000</b>		
Carry trade return $\nrightarrow$ Gold price	Copper price	2.940	0.172	0.506	<b>0.000</b>	0.963	0.221	<b>0.000</b>		

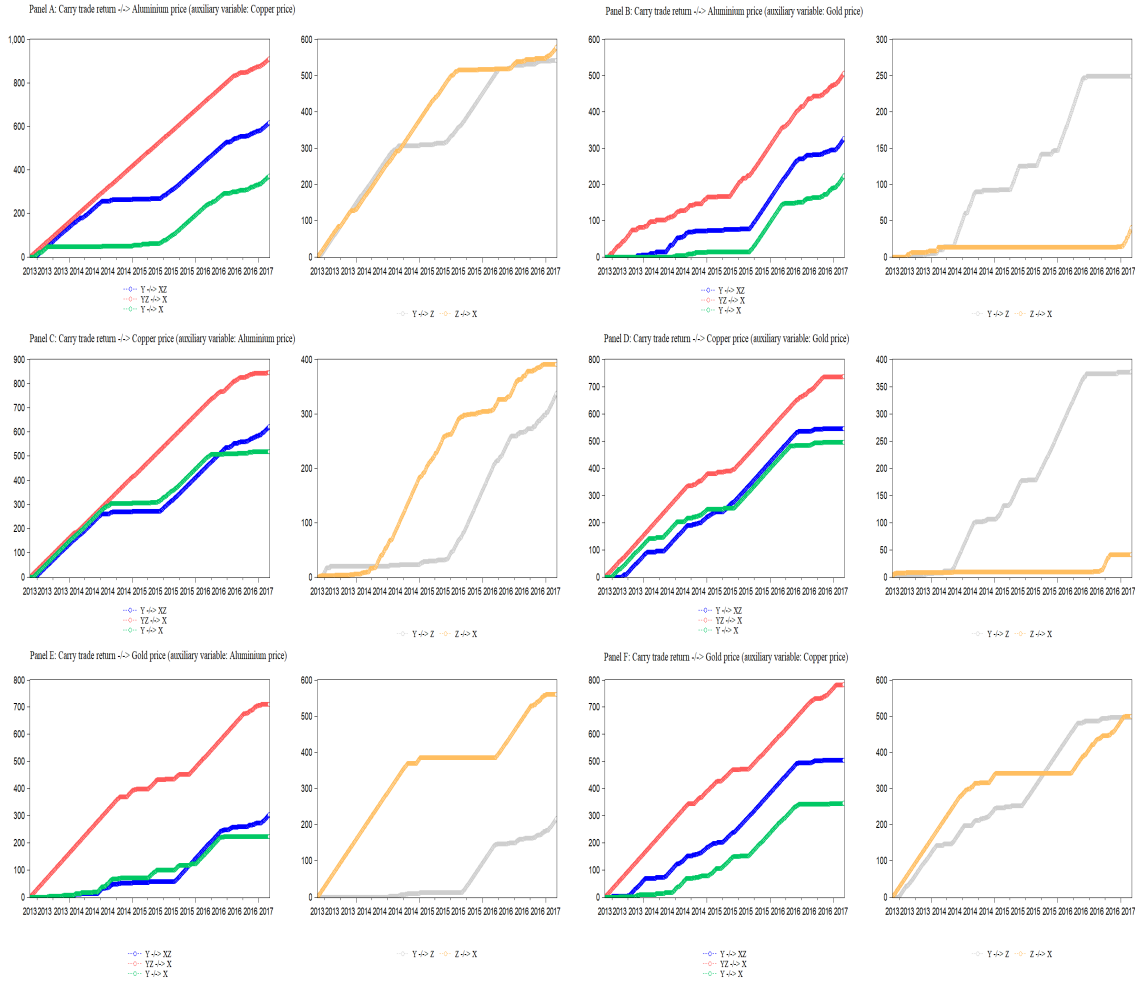
*Note:* The relationship  $y \nrightarrow x$  stands for  $y$  does not cause  $x$ . The size of the fixed rolling-window is 320 days. The maximum order of the VAR model is 8 lags. Avg.  $p$ -values are the average of bootstrap  $p$ -values. Bootstrap iterations are 1000 times. Bootstrap  $p$ -values of less than 5% indicate causality within that window. The average VAR order and average  $p$ -values are authors' calculation. Bolded types signify cases in which the null hypothesis of non-causality is rejected.

**Fig.5:** Rolling window  $p$ -values of hypothesis testing rejection frequencies (commodity price  $\leftrightarrow$  carry trade return)



*Note:* The relationship  $y \leftrightarrow x$  stands for  $y$  does not cause  $x$ . Rejection frequencies are generated based on the approximate  $p$ -values.

**Fig.6:** Rolling window  $p$ -values of hypothesis testing rejection frequencies (carry trade return  $\neq$  commodity price)



*Note:* The relationship  $y \not\rightarrow x$  stands for  $y$  does not cause  $x$ . Rejection frequencies are generated based on the approximate  $p$ -values.

of tests contradict each other. Furthermore, causality appears at any horizon if and only if it takes place at horizon one (first day in each window). The non-rejection of either test 0.1 or test 0.2 in both Panel A and B (rejection frequencies are highlighted in Fig.5 and 6) indicate that fluctuations in commodity prices never anticipate rise in carry trade returns, vice versa. When we proceed to check individual horizons, we find the causal effect continues to be of indirect nature in all the relationships between carry trade returns and commodity prices (rejection of test 1.0).

We find a causal chain through copper price (rejection of tests 1.1 and 1.2). The causal chain imposes the existence of causation delays in the response of aluminium price to changes in carry trade returns. Specifically, aluminium price responds with at least three days delay to carry trade returns changes. Furthermore, we obtain evidence on broken causal chains between commodity prices to returns of carry trade. For example, in the case of aluminium price, we find that change in aluminium price causes fluctuations in gold price (rejection of test 1.1), yet gold price does not cause carry trade return (non-rejection of test 1.2). Therefore, a causal linkage from aluminium price to carry trade returns cannot be established. Similarly, while copper price causes gold price, yet gold price does not in turn cause returns of carry trade, thus again a causal relationship from copper price to carry trade returns through gold price cannot be inferred.

For comparison, we discuss the causality results that would obtain if the [Toda and Yamamoto \(1995\)](#) approach has been used for causality analysis. The results in [Table A2](#) in Appendix outline a unidirectional causal relationship running from carry trade returns to copper price and a bidirectional causality between gold price and returns of carry trade. Hence, horizontal-specific causality tests are capable of revealing a causal chain between carry trade returns and aluminium price (transmitted through copper price).

## 7 Alternate linkage between carry trade and commodity prices

This section reports the results that address the robustness of the estimates presented above. We begin with an examination of the robustness of our results to the use of alternative way to conduct carry trade strategies (futures contract). [Section 7.1](#) establishes a revised [Frankel and Rose \(2010\)](#) model that considers the determinants of commodity prices under interest rate control regime. This is followed by estimating the model to the use of structural VAR (SVAR) in [Section 7.2](#). The results of SVAR is reported in [Section 7.3](#).

### 7.1 Theoretical Framework

Commodities have the characteristics of storage and relative homogeneity, therefore they have dual attributes of both assets and goods. For the former attributes, the supply and demand of inventory affects its prices. In terms of the latter attributes, the production and consumption of commodities decide the prices in inter-temporal period. In this section, starting from the asset attributes of commodities, we investigate the formation of its prices. Following [Frankel and Rose \(2010\)](#), the equation used to decide commodity prices can be derived from expectation formulation condition and conditions for arbitrage.

Let  $s$  denotes spot price,  $p$  represents inflation rate, then the real price of commodity is given

by  $q = s - p$ ,  $\bar{q}$  stands for the long-term equilibrium commodity prices. In the case of rational expectation, if investors observe the real commodity prices in the current period is higher or lower than its long run equilibrium, it is reasonable for them to expect the prices will return to the equilibrium in the future. Following Frankel (1986), assume the adjusting rate to long run equilibrium is  $\theta$  ( $\theta > 0$ ), which can be also specified as follow:

$$E[\Delta(s - p)] \equiv E(\Delta q) = -\theta(q - \bar{q}) \quad (8)$$

Rearrange the above equation, we can get:

$$E(\Delta s) = -\theta(q - \bar{q}) + E(\Delta p) \quad (9)$$

Considering the role of commodity futures market and futures price in expectations of commodity prices and production decision making, we introduce an extended extrapolation expectation formation mechanism. Specifically, the expected prices in the current period are the summation of the real prices and a certain proportion of the momentum of futures price (namely, smoothing coefficient  $\rho$ ) in the preceding period.<sup>14</sup> Let  $f$  represents forward or futures price, then Eq.(9) can be written as:

$$E(\Delta s) = -\theta(q - \bar{q}) + E(\Delta p) + \rho(\Delta f_{-1}) \quad (10)$$

At the same time, investors decide to hold commodities for another period, or to sell it at today's prices and use the proceeds to earn interest. The expected rate of return for the two alternatives should be equal, that is, satisfying the conditions for arbitrage. Thus, we have:

$$E(\Delta s) + c = i \quad (11)$$

where  $i$  is nominal interest rate,  $c$  denotes net benefit and  $c = a - b - d$ . The term  $a$  stands for convenience yield for holding commodities in stock, which used to deal with supply disruptions or convenience preferences;  $b$  represents storage cost of commodities; and  $d$  refers to risk premium. Furthermore,  $d = E(\Delta s) - (f - s)$ . Therefore, net benefit  $c$  is the benefit after considering the storage cost and risk premium.

We now using Eq.(10) to investigate the influence of futures price on spot price after commodity market liberalization. Combining Eq.(10) and Eq.(11), we can obtain:

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<sup>14</sup> In a typical extrapolation expectation model, the price in time  $t$  is the aggregate of the price in time  $t - 1$  and a certain proportion of the price difference between time  $t - 1$  and  $t - 2$ . However, owing to the role and existence of commodity futures market, lagged futures price contains more expected information. Therefore, when the futures price is available, it is better and more economical to replace the lagged spot price difference with the lagged futures price.

$$i - c = -\theta(q - \bar{q}) + E(\Delta p) + \rho(\Delta f_{-1})$$

Solving for  $(q - \bar{q})$ , can get:

$$q - \bar{q} = -\left(\frac{1}{\theta}\right)(i - E(\Delta p) - c) + \frac{\rho}{\theta}(\Delta f_{-1}) \quad (12)$$

It can be seen from Eq.(12) that the real commodity prices have a negative relationship with the difference of real interest rate and net benefit. Therefore, when real interest rate is high (low), to have expectation of rise (fall) in future commodity prices and satisfy conditions for arbitrage, money outflows (inflows) from commodity market until the commodity prices are lower (higher) than its long-run equilibrium values.

Plugging  $c = a - b - d$  into Eq.(12), can obtain:

$$q = \bar{q} - \left(\frac{1}{\theta}\right)(i - E(\Delta p)) + \frac{1}{\theta}a - \frac{1}{\theta}b - \frac{1}{\theta}c + \frac{\rho}{\theta}(\Delta f_{-1}) \quad (13)$$

Therefore, if the long term equilibrium commodity prices  $\bar{q}$  are given, there are factors other than real interest rate that still affect real commodity prices. These factors include net benefit, storage cost, risk premium and a certain proportion of the momentum of futures price in the preceding period. Next, we discuss the proxies used for these factors.

First, the state of economic activity is the main factor that determines the net benefit. Specifically, rising economic activity can stimulate demand of inventory due to net benefit, hence has a positive effect on commodity prices. We use the index of industrial product (IP) that can reflect the performance of an economy as a proxy for the net benefit.

Second, considering the short term stability of storage capacity, storage cost increases with the storage approaching its existing capacity. Assume  $b = \Phi(v)$ , where  $v$  stands for inventories. If the inventory level was observed at a historic high, then the cost of storage must be high. Thus, the inventory level has a negative impact on commodity prices. Consequently, the level of inventory is used as a proxy for storage cost.

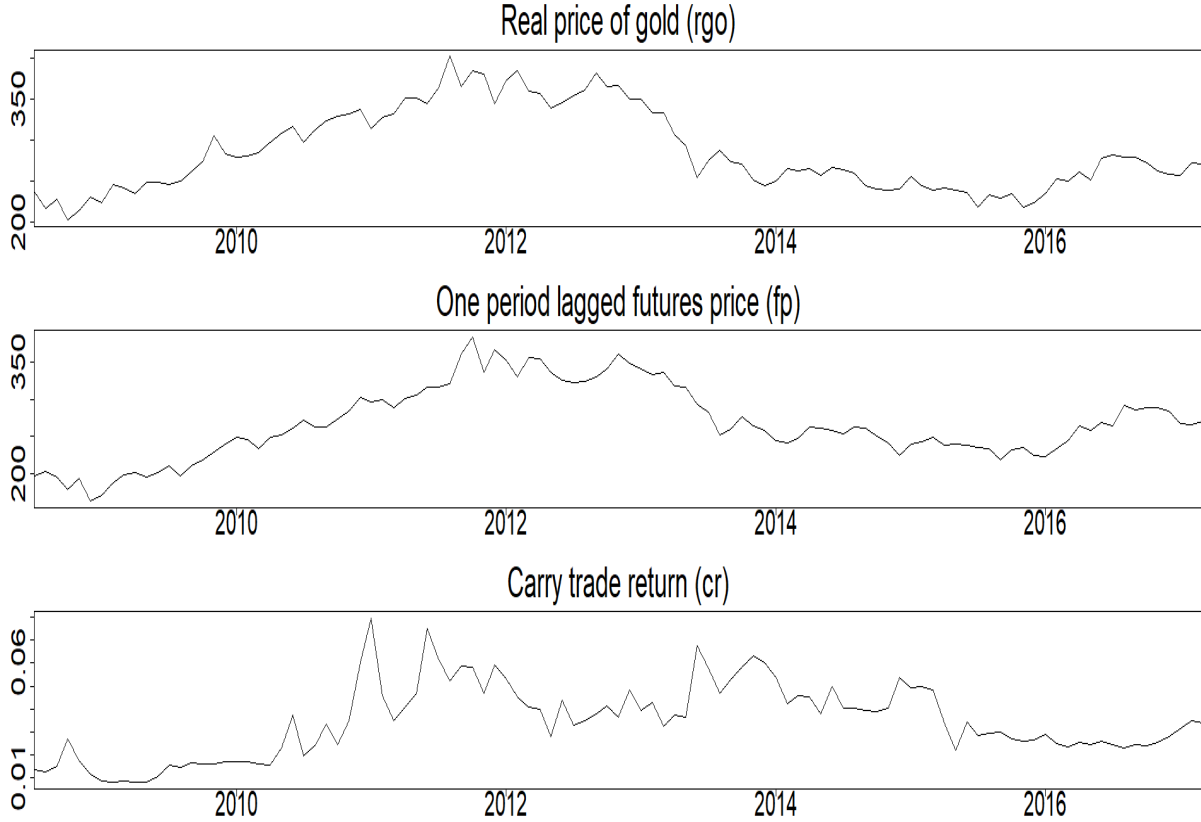
Third, we use commodity market volatility ( $\sigma$ ) difference of futures price and spot price ( $f - s$ ) as proxies for risk premium.

Substituting these proxies into Eq.(13), we can get:

$$q = C - \left(\frac{1}{\theta}\right)(i - E(\Delta p)) + \frac{1}{\theta}\gamma(y) - \frac{1}{\theta}\Phi(v) + \delta(f - s) + \frac{\rho}{\theta}(\Delta f_{-1}) \quad (14)$$

where  $C$  represents constant term and  $y$  refers to China's IP index.

**Fig.7:** Historical evolution of the series, 2008:07-2017:03



Considering the interest rate controls in China, interest rate variables can be replaced by monetary variables. To do so, we introduce a money demand function:  $\frac{M}{P} = m(y, i - E(\Delta p))$ , where demand for money is positively related to aggregate economic activities, but negatively related to real interest rate. Adding the function of economic activities on commodity prices ( $\alpha(y)$ ) and the function of economic activities on money demand to generate a new function  $\beta(y)$ , where  $y$  still has a positive relationship with commodity prices. The function of money supply on commodity prices can be specified as  $k(\frac{M}{P})$ . In addition, our variables of interest are futures price and carry trade returns. Thus, plugging the two functions and carry trade returns ( $CR$ )

into Eq.(14), can obtain:

$$q = C + \frac{1}{\theta}k\left(\frac{M}{P}\right) + \frac{1}{\theta}\alpha(y) - \frac{1}{\theta}\Phi(v) + \delta(f - s) + \frac{\rho}{\theta}(\Delta f_{-1}) + \beta CR \quad (15)$$

Eq.(15) is the empirical usable form of the model tested in Section 7.2. For simplicity, we assume that the relationship between variables is linear.

Fig.7 shows the historical evolution of our series of interest (real price of gold,  $rgo_t$ , one period lagged futures price,  $fp_t$ , and carry trade return,  $cr_t$ ) over the sample period.

## 7.2 Structural VAR identification

Using non structural method, Sims (1980) proposed a new approach to examine the relationship between variables: vector autoregression (VAR) model. Although VAR model has been widely used in commodity prices literature (e.g. Gutierrez1 and Piras, 2014; Kawamoto et al., 2011), there are several drawbacks of the framework: First, the VAR model contains too many parameters. Only VAR models with fewer variables can get satisfactory results through OLS and maximum likelihood estimations. Second, VAR model does not consider economic theory, the impulse response generated cannot be identified as intrinsic structural error due to innovation, hence no structural explanations can be provided. In this paper, we adopt SVAR model developed by Blanchard and Quah (1989) to disentangle the underlying causes of gold price fluctuations. The SVAR overcomes the above issues by imposing constraints on the parameter space to reduce the estimated parameters. This approach also enables us to measure the contribution of these components to observed gold price throughout the sample period. Furthermore, the structural shocks identified by the model enables us to estimate counterfactual prices in the absence of one or more of the components.

We include seven variables in a vector  $y_t$ : (i) real price of gold ( $rgo$ ), (ii) monetary liquidity ( $ml$ ), (iii) industrial production index ( $ip$ ), (iv) gold inventory ( $gi$ ), (v) risk premium ( $rp$ ), (vi) one period lagged futures price ( $fp$ ), (vii) carry trade return ( $cr$ ). The specification of our SVAR model is as follows:

$$A_0y_t = A_1y_{t-1} + A_2y_{t-2} + \dots + A_py_{t-p} + Cx_t + u_t \quad (16)$$

where  $x_t$  represents a vector of deterministic components that includes a constant and a linear trend,  $u_t$  stands for the structural shocks that are white noise and uncorrelated with each other.

We pre multiply Eq.(16) by  $A_0^{-1}$  to obtain the estimable reduced-form VAR:

$$A_0^{-1}Ay_t = \underbrace{A_0^{-1}A_1}_{G_1}y_{t-1} + \underbrace{A_0^{-1}A_2}_{G_2}y_{t-2} + \dots + \underbrace{A_0^{-1}A_p}_{G_p}y_{t-p} + \underbrace{A_0^{-1}C}_{G_0}x_t + \underbrace{A_0^{-1}u_t}_{\epsilon_t}$$

where  $A_0^{-1}A = I$ ,  $I$  is the identity matrix. Write the above equation in compact form, can get:

$$y_t = G_1y_{t-1} + G_2y_{t-2} + \dots + G_py_{t-p} + G_0x_t + \epsilon_t \quad (17)$$

where the reduced-form shocks,  $\epsilon_t$ , are prediction errors and are a weighted sum of the structural shocks, the matrix  $A_0$  provides those weights (i.e.  $\epsilon_t = A_0^{-1}u_t$ ). It requires making sufficient assumptions to enable consistent estimation of the unknown elements of  $A_0$  to identify the structural shocks.

We postulate that  $A_0^{-1}$  has a recursive structure such that the reduced-form errors  $\epsilon_t$  can be decomposed according to  $\epsilon_t = A_0^{-1}u_t$ :

$$\epsilon_t \equiv \begin{pmatrix} \epsilon_t^{rgo} \\ \epsilon_t^{ml} \\ \epsilon_t^{ip} \\ \epsilon_t^{gi} \\ \epsilon_t^{rp} \\ \epsilon_t^{fp} \\ \epsilon_t^{cr} \end{pmatrix} = \begin{bmatrix} 1 & 0 & 0 & 0 & 0 & 0 & 0 \\ a_{21} & 1 & 0 & 0 & 0 & 0 & 0 \\ a_{31} & a_{32} & 1 & 0 & 0 & 0 & 0 \\ a_{41} & a_{42} & a_{43} & 1 & 0 & 0 & 0 \\ a_{51} & a_{52} & a_{53} & a_{54} & 1 & 0 & 0 \\ a_{61} & a_{62} & a_{63} & a_{64} & a_{65} & 1 & 0 \\ a_{71} & a_{72} & a_{73} & a_{74} & a_{75} & a_{76} & 1 \end{bmatrix} \begin{pmatrix} u_t^{rgo} \\ u_t^{ml} \\ u_t^{ip} \\ u_t^{gi} \\ u_t^{rp} \\ u_t^{fp} \\ u_t^{cr} \end{pmatrix} \quad (18)$$

### 7.3 Results

Panel B of [Table 2](#) reports the stationarity results. Majority of the tests suggest that all the variables are nonstationary in levels. The results of unit root testing using the first difference of the variables that are non-stationary in levels indicate that except the real price of gold, the other variables are all first difference stationary at the 5% level of significance or better. The ADF and PP tests provide conflicting results regarding the stationarity of real price of gold in first difference. As argued in [Myers \(1992\)](#), compared with the ADF test, PP test is more robust to autocorrelation and heteroskedasticity in the errors, two issues which plague high frequency time series data. The PP test statistic reveals that the first differenced real price of gold is stationary at the 1% level of significance. One common feature of high frequency commodity price series (data sampled at daily, weekly or monthly intervals) is that they appear to contain

trends which change randomly over time. One of the problems with the ADF and PP tests is that the null of stochastic trend. This ensures that a stochastic trend cannot be rejected unless there is strong evidence against it. A stochastic trend may not be rejected simply because the data are not informative about whether or not there is a unit root in the series. In response, KPSS test examines the null of stationarity against the alternative hypothesis that the series has a stochastic trend. The test statistic for KPSS test confirms that the real price of gold is stationary in first difference.<sup>15</sup> Overall, we conclude that all the variables are  $I(1)$ .

We then conduct [Johansen \(1988, 1991\)](#) cointegration test among variables. Cointegration among variables indicates that there exists an economic mechanism that restricts the relationship between variables, so that the deviation between variables is limited in the short run, then gradually achieve equilibrium in the long term. When there is cointegration among variables, the SVAR model is stable so that the impulse response function and variance decomposition are meaningful. [Table 9](#) presents the cointegration test results. Both trace statistic and maximum eigenvalue statistic imply that there is one cointegrating equation among variables, meaning that all variables move together in the long run, thus SVAR model can be established.

**Table 9:** [Johansen \(1988, 1991\)](#) cointegration test results

No. of cointegration equations	Trace Statistic	<a href="#">MacKinnon-Haug-Michelis (1999)</a> $p$ -values	Max-Eigen Statistic	<a href="#">MacKinnon-Haug-Michelis (1999)</a> $p$ -values
None	160.795***	0.000	68.307***	0.000
At most 1	92.488	0.082	32.130	0.296

*Note:* Lag order selected is four. \*\* Denotes statistically significant at the 5% level.

**Table 10:** Results of coefficient estimates

Variables of interest	Coefficient	Std. error	$z$ -statistic	$p$ -values
<b>One period lagged futures price</b>	0.341	0.080	4.267	0.000
<b>Carry trade returns</b>	0.582	0.299	1.950	0.051
Monetary liquidity	0.048	0.113	0.427	0.669
IP index	0.022	0.096	0.229	0.819
Gold inventory	0.240	0.125	1.912	0.056
Risk premium	0.986	0.038	25.875	0.000

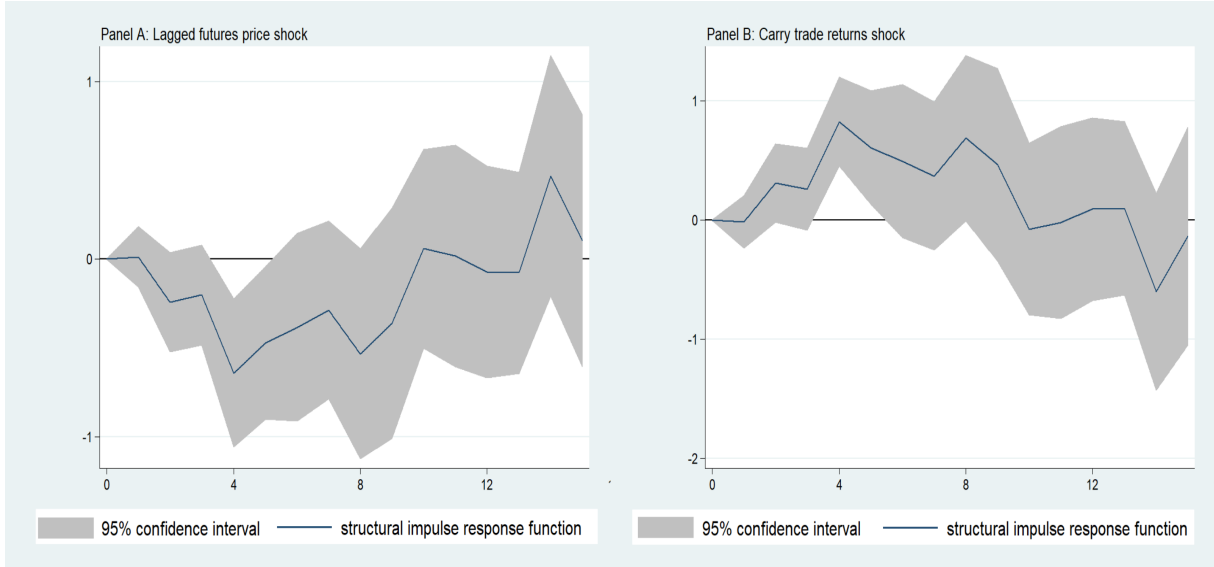
*Note:* Text in bold stands for variables of interest.

[Table 10](#) reports the coefficient estimates of all the variables. Both futures price in last period

<sup>15</sup> It is certainly a plausible hypothesis that results of some traditional unit root tests have been affected by errors in the test specification relative to the true data generating process. For example, the ADF test result is biased toward a false acceptance of a unit root in the presence of structural breaks. We, therefore, use [Narayan et al. \(2016\)](#) test to check the robustness of stationarity results for real price of gold. The test statistics identify the real price of gold is  $I(1)$ . Results of the test are available upon request.

and carry trade returns have a positive significant effect on real price of gold. Our findings imply that real price of gold is driven by commodity trade financing and returns of carry trade.

**Fig.8:** Responses of the real price of gold to structural shocks



**Table 11:** Percent contribution of lagged futures price and carry trade returns shocks to the overall variability of the real price of gold

Horizon	Shocks	
	Lagged futures price	Carry trade returns
1	0.000	0.000
2	2.461	40.327
3	4.246	71.554
15	3.174	53.641
$\infty$	3.539	59.725

*Note:* Based on variance decomposition of the SVAR model (16).

Fig.8 reviews the responses of the real price of gold to the lagged futures price shock and the carry trade returns shock. Point estimates are indicated by the blue solid line and the shaded grey areas around the blue lines delineate the 95% confidence intervals. As can be seen from Fig.8, the two structural shocks have very different impacts on the real price of gold. Specifically, an unexpected increase in the futures price in last period causes an immediate and sustained decrease

in the real price of gold within the first four months, then the impact gradually disappears. An unexpected rise in carry trade returns has a transitory positive effect within the first eight months, before declining below baseline in subsequent periods.

The forecast-error-variance decompositions in [Table 11](#) quantify the effects of the structural shocks on the real price of gold. Although in the short-run, the impacts of the two structural shocks on real gold prices are zero, the explanatory power increases as the forecast horizon increases. In the long run, the returns of carry trade shock accounts for about 60% of the variability in the real price of gold. This indicates that structural shock in the currency market is an important fundamental for the gold price in China.

## 8 Concluding Remarks

This study investigates the relationship between carry trade returns and prices of commodities (aluminium, copper and gold) which are most commonly used as collateral assets in carry trade deals. We focus on the U.S. dollar -Chinese RMB exchange rate, where China (target country for the carry trade deals) is one of the world's fastest growing economies with rigorous capital and interest rate control regimes and the U.S. represents funding country with low interest rates post GFC. In particular, we analyse the existence of the long run equilibrium relationship between carry trade returns and collateral prices using the ARDL model with structural breaks. We also study causality between the carry trade returns and prices of collateral assets using [Hill \(2007\)](#) sequential causality tests. Lastly, in line with [Frankel and Rose \(2010\)](#) we employ the SVAR framework to estimate an alternative model of commodity prices determination under the interest rate controls.

Our results indicate that in the long run copper price has a positive impact on carry trade returns. By contrast, there is a negative relationship between gold price and returns of carry trade. These could be because the fact that copper and gold are the two most commonly used commodity collateral to conduct carry trade strategies due to the nature of these commodities. For example, gold can be used as investors as a safe heaven asset in times of uncertainty as suggested by [Baur and Lucey \(2010\)](#), [Baur and McDermott \(2010\)](#) and [Beckmann et al. \(2015\)](#). Copper can be viewed as the important indicator of a country's economic activity that mirrors business cycles. In particular, being an industrial nonferrous metal, copper has an important role in industrial manufacturing. Changes in the copper price volatility have impact on industrial production, policy decisions by governments as well as the risk management plans and portfolio allocation decisions by investors and traders (see [Todorova et al., 2014](#); [Gong and Lin, 2018](#)). In addition, the tiny positive effect of copper price and large negative effect of gold price imply that

there are hedge characteristics for copper and gold returns on returns of carry trade in the long term.

In the short run, however, we did not find any evidence of association between the prices of aluminium and copper and carry trade returns, while we observe a mix of positive and negative impact of gold price on the carry trade returns as well. We also found significant structural breaks corresponding to major economic shocks affecting the long-run path of the carry trade returns and commodity prices. In particular, we found that GFC badly affected Chinese carry traders as capital started moving away from China since early 2008. This happened because of liquidity shortage, unwinding of carry trades and flight to safety of the U.S. government stocks by international investors (Yu, 2010). The real impact of the carry trade unwind is hard to measure. Following the IMF,<sup>16</sup> due to the difficulties with recording data on carry trade positions, the dynamics of capital flows are driven by the loans and deposits part of Other Investment section of the China's Capital Account on the Balance of Payments, which are "two relatively open parts of the capital account in which the carry trade was played". At the same time, carry trade positions can end up in Errors and Omissions section of the Chinese Balance of Payments. It should also be noted that judging based on the net capital inflows post GFC, there was a sharp unwind of the carry trade positions which was however short lived and in 2009, capital started going back to China. This finding of the carry trade unwind due to the financial crisis confirms empirical evidence on other carry trade pairs such as for example Australian dollar - Japanese Yen as shown by Kim (2016) and Reserve Bank of Australia.<sup>17</sup>

The results of Hill (2007) sequential causality testing suggest that multi-horizon causality testing does uncover crucial information on the dynamic interaction among carry trade returns and commodity prices of the three collateral assets chosen for the analysis and confirm that structural breaks, if they exist, appear to be crucial for causality inference. In regard to causality direction, we find a causal chain through copper price. In particular, it takes at least three days for carry trade returns to influence aluminium price. Furthermore, there are broken causal chains between prices of aluminium and copper to carry trade returns (transmitted via gold price).

We develop a revised Frankel and Rose (2010) model to investigate the determinants of the real gold price under the interest rate control regime. In doing so, we focus on one period lagged futures price shock and carry trade returns shock. Our results, based on an identified SVAR, show that the response of the real price of gold may differ greatly depending on various shocks. In particular, we show that shocks in the currency market account for about 60% of the long run

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<sup>16</sup> Refer the IMF Country Report No.16/271, available at: <https://www.imf.org/external/pubs/ft/scr/2016/cr16271.pdf>

<sup>17</sup> Refer the article "Japanese retail investors and the carry trade", available on the website of Reserve Bank of Australia at: <https://www.rba.gov.au/publications/bulletin/2009/mar/pdf/bu-0309-1.pdf>

variability of the real price of gold.

There are several avenues for future research. First, this study assumes there is a linear relationship between carry trade returns and commodity prices. However, the relationship can be nonlinear for several reasons including the nature of data (i.e., negatively skewed returns and presence of nonlinear unit root), asymmetric risk profiles of market participants, asymmetric shipping costs and so on which could shed further light on this topic. Second, we considered only one exchange rate pair, while there are other carry trade currencies such as Australian dollar - Japanese Yen, U.S. dollar - Brazilian Real, etc for which no such studies exist. Third, we have chosen China as the special case due to the capital and interest rate control regimes introduced by the central bank. But other countries, e.g. Brazil, have also introduced capital controls. Studying a panel of countries which introduced capital controls would give a richer picture on the impact of carry trade returns and prices of collateral assets since collateral assets are used to bypass capital controls. Fourth, the focus of this analysis were prices of the commodity-based collateral assets. In this analysis we did not consider other forms of collateral. For example, shares are traditionally used as collateral in Japan and other countries in Asia. Analyzing other different forms of collateral assets would greatly contribute to the understanding the linkages between carry trade returns and prices of collateral assets.

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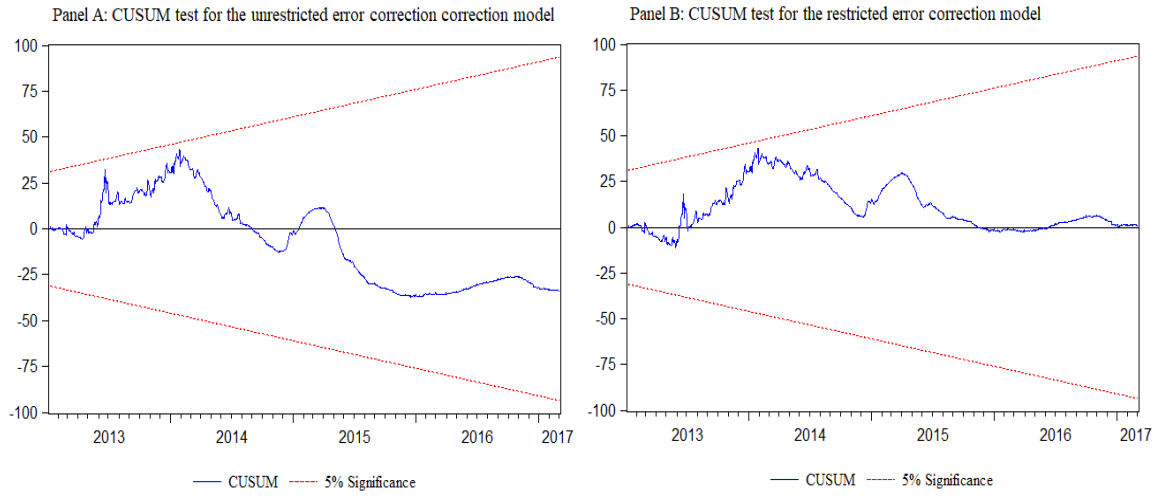
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# Appendix

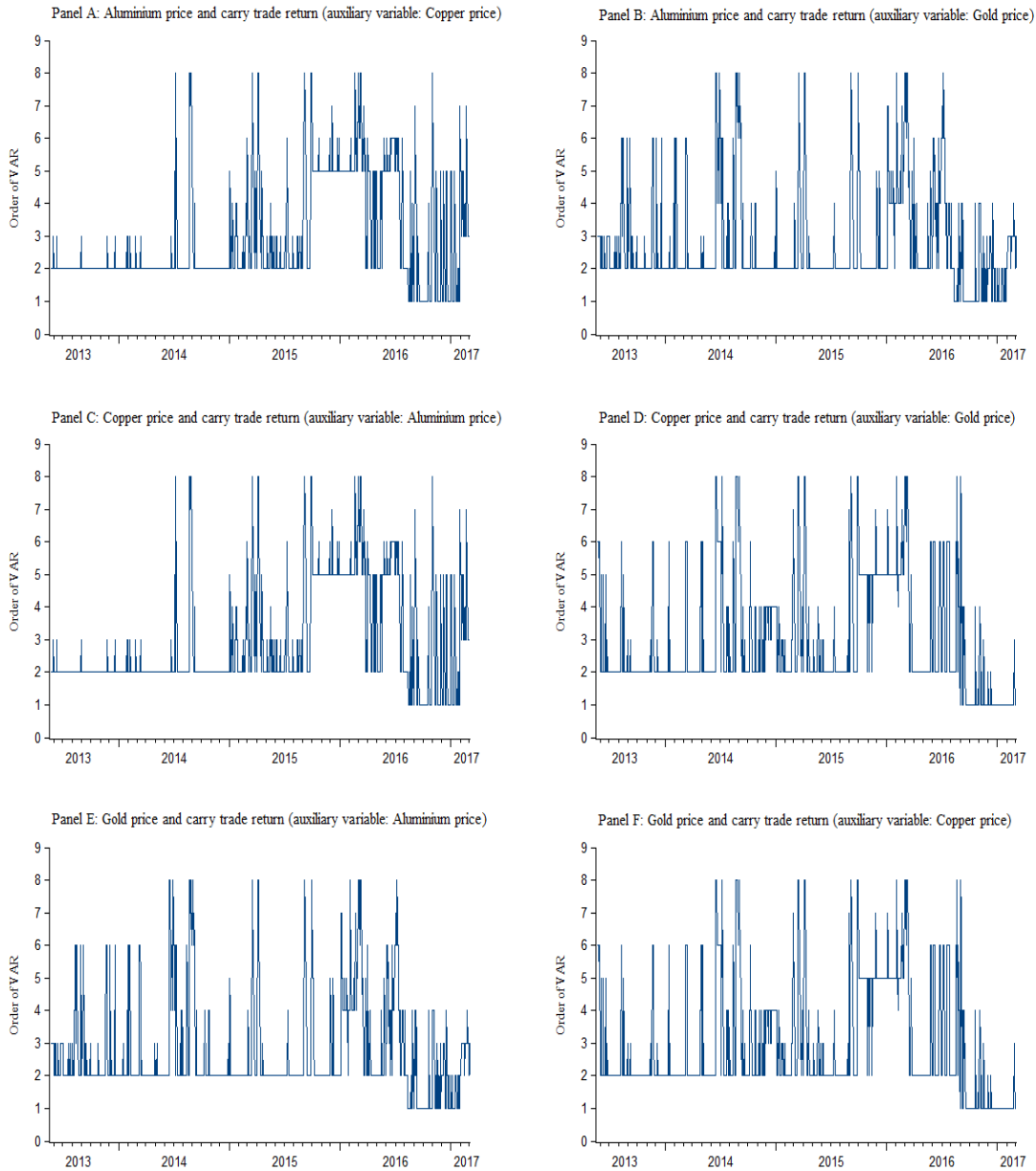
**Table A1:** Detail of data resource and definition

No.	Variable	Component	Source
<b>Panel A: Capital control</b>			
1	Aluminium Price Premium	Aluminium price premium in Shanghai Bonded Warehouse (USD per Metric Tonne)	Thomson Reuters Datastream
2	Aluminium Cash Price	Aluminium Cash Price in London Metal Exchange (LME) (USD per Metric Tonne)	Thomson Reuters Datastream
3	Aluminium Price	Aluminium Price Premium + Aluminium Cash Price	Authors' Calculation
4	Copper Price Premium	Copper price premium in Shanghai Bonded Warehouse (USD per Metric Tonne)	Thomson Reuters Datastream
5	Copper Cash Price	Copper Cash Price in London Metal Exchange (LME) (USD per Metric Tonne)	Thomson Reuters Datastream
6	Copper Price	Copper Price Premium + Copper Cash Price	Authors' Calculation
7	Gold Price	Gold (acceptable purities: 99.99) Close Price (USD per Gramme) in Shanghai Gold Exchange (SGE)	Thomson Reuters Datastream
8	Onshore risk-free Interest Rate	Shanghai 1 month Interbank Offered Rate (in basis point)	Thomson Reuters Datastream
9	Offshore risk-free Interest Rate	Federal Funds Rate (in basis point)	Thomson Reuters Datastream
10	Foreign Exchange Spot Rate	CNY to USD Exchange Rate (in basis point)	Thomson Reuters Datastream
11	Foreign Exchange Forward Rate	CNY to USD 3-month Forward Rate (in basis point)	Thomson Reuters Datastream
12	Carry Trade Return	Rate of return on carry trade obtained using <a href="#">Eq.(1)</a>	Authors' Calculation
<b>Panel B: Interest rate control</b>			
13	Gold Spot Price	Gold (acceptable purities: 99.99) Close Price (CNY per Gramme) in SGE	Thomson Reuters Datastream
14	Real Price of Gold	Gold spot price deflated by CPI	Authors' Calculation
15	SHFE Gold Futures Price	Continuous trading settlement gold price (CNY per Gramme) in Shanghai Futures Exchange (SHFE) (contract size: 1 kilogram/lot)	Thomson Reuters Datastream
16	NYMEX Gold Futures Price	Continuous trading settlement gold price (CNY per Gramme) in New York Mercantile Exchange (NYMEX) (contract size: 100 troy ounce)	Authors' Calculation based on gold futures price (USD per troy ounce) in NYMEX
17	Industrial Production Index	China's Industrial Production Index	Thomson Reuters Datastream
18	Monetary Liquidity	(i) China's broad money (M2) supply (hundred million yuan) (ii) China's GDP (hundred million yuan)	China's M2: Thomson Reuters Datastream China's GDP: Author's calculation on converting quarterly GDP to monthly GDP based on monthly change rate of IPI
19	Gold Inventory	Total gold stock of the warehouse (on warrant)	Thomson Reuters Datastream (we then convert unit from tons to gramme)
20	Risk Premium	Gold futures price in NYMEX - Gold spot price in SGE	Author's calculation
21	Foreign Exchange Spot Rate	CNY to USD Monthly Exchange Rate	Thomson Reuters Datastream
22	Carry Trade Return	Monthly rate of return on carry trade obtained using <a href="#">Eq.(1)</a>	Authors' Calculation

**Fig.A1:** Cusum test for the ARDL framework



**Fig.A2:** Rolling window VAR orders ( $p$ )



*Note:* Each panel contains two trivariate VAR equations with the same auxiliary variable. For example, the two trivariate VAR models included in Panel A are: Aluminum price  $\nrightarrow$  carry trade return (auxiliary variable: Copper price), Carry trade return  $\nrightarrow$  Aluminum price (auxiliary variable: Copper price). The relationship  $y \nrightarrow x$  stands for  $y$  does not cause  $x$ .

**Table A2:** Toda and Yamamoto (1995) causality test

Causality Direction	No. of lags	Test statistic	<i>p</i> -values
<b>Panel A:</b> Commodity prices to carry trade returns			
Aluminium price -/->Carry trade return	3	0.468	0.791
Copper price -/->Carry trade return	10	2.594	0.978
Gold price -/->Carry trade return	13	19.758	0.072
<b>Panel B:</b> Carry trade returns to commodity prices			
Carry trade return -/->Aluminium price	3	1.033	0.597
Carry trade return -/->Copper price	10	15.056	0.089
Carry trade return -/->Gold price	13	23.242	0.026

*Note:* The relationship  $y \text{ -/-> } x$  stands for  $y$  does not cause  $x$ .