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Abstract

This paper investigates whether commodity convenience yields - the yields that accrue to the holders of physical commodities - can predict the exchange rate of commodity-exporters' currencies. Predictability is a consequence of the fact that i) convenience yields are useful predictors for commodity prices and ii) commodity currencies have a strong relationship with commodity prices. The empirical evidence indicates that there is a significant relationship between aggregate measures of convenience yields and commodity currencies' exchange rate, both in-sample and out-of-sample. A high level of convenience yields strongly predicts a depreciation of the Australian, Canadian and New Zealand dollars exchange rates at horizons of 1 to 24 months.

Keywords: Exchange rates, Commodity currencies, Convenience yield

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1 Introduction

The Australian, Canadian and New Zealand dollars have in common that they are currencies of net primary commodity exporting countries and have a long history of a floating exchange rate. As a consequence of these features, their nominal and real exchange rates are directly affected by fluctuations in the world price of commodities. These attributes have been exploited to identify the effect of terms-of-trade shocks on the real exchange rate (Amano and Van Norden, 1995; Chen and Rogoff, 2003) or test the asset market approach of exchange rate determination (Chen et al., 2010). Commodity currencies - as these currencies are sometimes termed in the literature - are considered by many researchers as an exception to the well-documented Meese-Rogoff puzzle, which is that the current exchange rate is often a better predictor of future exchange rates than a linear combination of macroeconomic fundamentals (see Meese and Rogoff, 1983; Chen and Rogoff, 2012). Although the addition of contemporaneous commodity price changes to standard macroeconomic fundamentals generally improves the in-sample fit of empirical nominal exchange rate models, the evidence on out-of-sample forecasting performance is mixed.

This paper proposes and evaluates a novel approach to forecast the exchange rate of commodity currencies using commodity convenience yields, which are defined as the yields that accrue to the holder of commodity inventories. Convenience yields might be useful for predicting commodity currencies because they predict future changes in commodity prices and there is a strong relationship between commodity prices and commodity currencies' exchange rates. Empirically, I find that future changes in the bilateral exchange rates of the Australian, the Canadian and New Zealand Dollars vis a vis the US Dollar or the UK Pound are significantly related to commodity convenience yields. A high level of convenience yields predicts a depreciation of all three exchange rates in horizons of 12 to 24 months and also at shorter horizons for the Australian and New Zealand dollars. Exchange rate forecasts based on convenience yields outperform the random walk model and also outperform forecasts based on contemporaneous commodity price changes.

In the theory of commodity storage, the convenience yield is a benefit that accrues to the holder of an inventory. Commodity inventories have an option value as they allow a producer using the commodity as an input to meet unexpected demand for his produced good. They also allow to overcome situations of aggregate stock-out or disruptions in the supply chain. Gorton et al. (2012) empirically document the convenience yield of 31 commodities and find that it is a decreasing, non-linear function of inventories. A high level of convenience yield, due to a low level of aggregate inventories, precedes a decrease in the price of the commodity as inventories return to their normal level. In the model of rational commodity pricing of Pindyck (1993) commodity convenience yields are forward-looking variables that incorporate information about future supply and demand conditions.

My empirical investigation relies on an aggregate measure of commodity convenience yields based on individual convenience yields. I calculate convenience yields of 21 commodities covering different commodity groups using spot and futures prices from April 1983 to January 2012. Then I perform a principal component analysis to extract common factors from the panel of convenience yields. The first principal component of convenience yield has strong predictive power for commodity currencies' exchange rate changes even after controlling for commodity price fluctuations.

A useful empirical characteristic of commodity convenience yields for forecasting commodity currencies is that they are persistent but stationary variables, whereas commodity prices are generally non-stationary. Engel et al. (2010) show that in present-value asset pricing models with discount factors close to unity, asset prices behave like a random walk and short-horizon regressions on fundamentals display low R-squared. This is the observed behavior of exchange rates and macroeconomic fundamentals. However in the presence of stationary fundamentals, long-horizon regression can have substantial power, even when the discount factor is close to one and the power of short-horizon is low. I therefore focus on different horizons from one month ahead changes in the exchange rate to 24-month ahead changes.

1.1 Related literature

This paper is related to the empirical literature focusing on the currencies of large commodity exporters to assess exchange rate models. Amano and Van Norden (1995) document a strong and robust relationship between the Canada - U.S. real exchange rate and terms of trade, proxied by the price of exported commodities relative to the price of imported manufactured goods. Chen and Rogoff (2003) find that the prices of commodity exports of Australia and New Zealand have a strong and stable influence on their real exchange rate, while Chen (2004) finds that incorporating commodity export prices into standard exchange rate models can generate a marked improvement in their in-sample performance. Chen et al. (2010) find evidence of Granger-causality from commodity prices to the exchange rate of commodity currencies which however does not translate into significant out-of-sample forecast ability. Closest to my paper is the recent work of Ferraro et al. (2011) who find some evidence of very short-term predictability of the Canadian / U.S. dollar nominal exchange rate using oil prices at the daily frequency.

Commodity prices have in common with exchange rates that they are difficult to predict and behave like random walks. Alquist et al. (2011) propose a comprehensive survey of the literature on forecasting oil prices. On the approach of using convenience yields for a predictive purpose Knetsch (2007) evaluates the use of oil convenience yields to forecast the price of oil. Finally, this paper has been inspired by a recent study of Gospodinov and Ng (2011) who show that the two leading principal components of commodity convenience yields have important predictive power for inflation, as they capture inflationary pressures.

2 Conceptual framework

2.1 Exchange rates and commodity prices

This paper considers the relationship between the nominal exchange rate of commodity exporters like Australia, Canada and New Zealand and world commodity prices. For these countries with a high share of exports in the commodity sector, commodity price fluctuations represent significant terms-of-trade shocks which may affect their floating exchange rates through different channels. Chen and Rogoff (2003) present a small open economy model with traded and non-traded goods and flexible labour markets to emphasize one potential channel. They show that an exogenous increase in the world price of a country's commodity exports has a positive impact on its real exchange rate through a channel similar to the effect of productivity shocks in a standard Balassa-Samuelson framework: wages and the demand for non-traded goods increase, exerting upward pressure on the price of non-traded goods. If the latter is not fully flexible, some of the adjustment to restore the efficient relative price between traded and non-traded goods has to be borne by the exchange rate, which will appreciate in response to a positive commodity price shock.

A second channel emphasized in the literature operates through the asset markets and a portfolio channel (see Chen and Rogoff, 2003; Chen, 2004; Chen et al., 2010; Chan et al., 2011). For an economy with a high share of exports in commodities, an exogenous increase in a country's exports price typically results in an improvement of the balance-of-payments and an accumulation of international reserves. These two factors lead to an increase in the relative demand for the country's currency leading to its appreciation.

This paper does not aim either at identifying one specific channel through which commodity prices affect the nominal exchange rate of commodity exporters nor at estimating structural parameters and therefore I consider a general reduced-form model for the (log-) exchange rate

$$e_t = a + b' f_t + E_t e_{t+1} \tag{1}$$

where e_t is the (log-) bilateral exchange rate, denominated as the foreign currency price of one unit of home currency, b a vector of reduced-form parameters, f_t a vector of exchange rate fundamentals of which p_t^{com}, the (log-) world commodity price, is one element and $E_t e_{t+1}$ is the expected future value of the exchange rate. Depending on the structural model considered, f_t could contain the differential between the home and foreign country of price level, industrial production, money supply or interest rates. A recent literature uses Taylor rules fundamentals as nominal exchange rate determinants (see Engel and West, 2005; Molodtsova and Papell, 2009).

The difference equation characterizing the exchange rate (1) can be solved forward and under the assumption that the first-differences of p_t and f_t follow first-order autoregressive processes the following expression for the first-difference of the exchange rate is obtained

$$\Delta e_t = \alpha + \beta' \Delta f_t + \epsilon_t \tag{2}$$

where the parameters α and β are functions of the parameters in equation (1).

This approach suggests that a way to forecast commodity currency returns is to use a forecast of the change in commodity prices. The next section outlines a strategy that has been proposed recently in the literature and which uses commodity convenience yields to obtain commodity price forecasts. Knetsch (2007) uses convenience yields to forecast the price of crude oil whereas Gospodinov and Ng (2011) show evidence that the two leading principal components of commodity convenience yields help predict commodity prices and also inflation, because the principal component capture inflationary pressures of the commodity prices.

2.2 Commodity prices and convenience yield

The framework to understand the forward-looking behavior of convenience yields for commodity prices is given by the theory of storage (see Kaldor, 1939; Brennan, 1958) which emphasizes the role of competitive inventory holders for linking commodity prices intertemporally. The convenience yield is an implicit benefit which accrues exclusively to the holder of a physical commodity. It has been introduced to explain situations in which positive inventories are held despite the fact that buying and holding the commodity is more costly than buying the commodity forward. Fama and French (1987) argue that the convenience yield arises because the commodity (eg. wheat) is an input in the production of other commodities (eg. flour). Inventories also help meeting unexpected demand and have an option value due to the positive probability of a stock-out which would imply additional production costs. Gorton et al. (2012) empirically document the convenience yield of 31 commodities and find that it is a decreasing, non-linear function of inventories. They further find that the relationship is affected by the storability of the commodity. It is weaker for commodities that are easy to store such as industrial metals and stronger for energies or agricultural commodities with strong seasonal factors.

Let P_t^j denote the spot price of commodity j, $F_{t,T}^j$ the price of a futures contract on commodity j with delivery at time T and $\varphi_{t,T}^j$ the net convenience yield (net of storage costs) that accrues to the holder of inventories of commodity j from time t to T. $\varphi_{t,T}^j$ is positive (negative) when the benefit of having the commodity in stock is higher (lower) than the storage costs (warehousing and insurance costs). Buying one marginal unit of commodity j at time t until T yields a payoff of $\varphi_{t,T}^j - (1 + i_{t,T})P_t^j$, with $i_{t,T}P_t^j$ being the (nominal) interest foregone from investing in the commodity from time t to T. Competitive storers adjust their inventory holdings to eliminate arbitrage opportunities. This occurs when the following no-arbitrage condition is satisfied

$$F_{t,T}^j = (1 + i_{t,T})P_t^j - \varphi_{t,T}^j \tag{3}$$

Although it is unobserved, the net marginal convenience yield can be measured from (3) as the (interest-adjusted) basis $(1 + i_{t,T})P_t^j - F_{t,T}^j$.

Another view on the relationship between commodity spot and futures prices is the theory of normal backwardation (see Keynes, 1930; Hicks, 1939), which emphasizes the risk premium earned by risk-averse investors for the uncertainty on future spot prices. According to this theory, current futures prices for delivery at time T are set at a discount of the expected future price. The size of the discount is given by the risk premium $\sigma^j_{t,T}$ that investors require to take long positions on the futures markets.

$$F^j_{t,T} = E_t[P^j_T] - \sigma^j_{t,T} \tag{4}$$

The combination of both theories summarized by (3) and (4) characterizes the forward-looking behavior of commodity prices with the following expectational difference equation:

$$E_t[P^j_T] - P^j_t = i_{t,T} P^j_t - \varphi^j_{t,T} + \sigma^j_{t,T} \tag{5}$$

which for a horizon of one period and an aggregate measure of commodity prices reduces to

$$E_t \Delta p^{com}_{t+1} = i_t p^{com}_t - \varphi_t + \sigma_t \tag{6}$$

The interest costs $i_t p^{com}_t$, the convenience yield φ_t and the risk premium σ_t are the three components of the expected change in the commodity prices.

3 Empirical methodology

3.1 Aggregate measures of convenience yields

The conceptual framework presented in Section 2 uses an aggregate measure of convenience yields without specifying how it is constructed. This section proposes two different measures computed from the convenience yields of individual commodities.

The focus on the effect of commodity prices on exchange rates through commodity exports in the conceptual framework suggests using an export-weighted country-specific average of the cross-section of individual commodity convenience as an aggregate measure of convenience yields. This implies the following measure

$$\overline{\varphi}_t = \frac{1}{N}\sum_{j=1}^{N} \omega^j \varphi_t^j \tag{7}$$

where ω^j is the share of commodity j in a country's total commodity exports and φ_t^j is the convenience yields of commodity j. The country-specific weights are reported in Table A2 in the Appendix.

A drawback of this measure is that a large share of commodities exported do not have futures markets with a long enough price history and are not included in the dataset. The exports-weighted average thus only covers a fraction of a country's commodity exports. On the other hand, some commodities not exported by the countries considered in this paper have had futures markets for a long time. Under the assumption that there are aggregate factors underlying the convenience yields of individual commodities, e.g. global demand shocks, incorporating those commodities into the analysis provides additional information about aggregate convenience yields. The strategy I follow is to extract principal components from a panel of individual convenience yields as in Gospodinov and Ng (2011). Principal components are weighted averages of the underlying individual series constructed to best explain the variation in the data. Formally, a principal component is an eigenvector corresponding to an eigenvalue of the $J \times J$ matrix $(NT)^{-1}\varphi_J'\varphi_J$, where φ_J is the $T \times J$ matrix that contains the T observations of the convenience yields of J commodities. Principal components are ordered according to their capacity to explain the variation in the data. I will use the first two principal components as predictors in the exchange rate model, while checking for the predictive ability of the next important principal components.

3.2 In-sample predictive ability of convenience yields

The econometric approach used to predict exchange rate returns derives from the conceptual framework. The exchange rate of a commodity currency is correlated contemporaneously with the price of the commodities exported by the country as shown in equation (2). Moreover the convenience yield of a commodity is a determinant of commodity price returns as shown by (6). The strategy used in this paper is to produce forecasts of exchange rate returns directly using commodity convenience yields and the relationship between exchange rates and commodity prices.

I use a standard regression framework to assess the in-sample predictive ability of convenience yields for exchange rates. I estimate the following empirical model that links exchange rate returns directly to commodity convenience yields

$$e_{t+h} - e_t = c^{(h)} + \alpha^{(h)}(L)\Delta e_t + \beta^{(h)}\varphi_t + \gamma^{(h)}X_t + \epsilon_{t+h}^{(h)} \tag{8}$$

where the dependent variable $e_{t+h} - e_t$ is the h-period change in the log exchange rate defined as the price in foreign currency of one unit of domestic currency, φ_t is one of the aggregate measure of commodity convenience yields described above and $\beta^{(h)}$ is the coefficient on which inference will be drawn to assess the predictive ability of convenience yields. I include lagged first differences of the exchange rate Δe_t to account for persistence in the exchange rate changes. I also include other determinants of the exchange rate in the vector X_t as the objective of the paper is to assess the incremental predictive ability of convenience yields beyond conventional predictors. In particular, I will include aggregate measures of commodity prices to test whether commodity convenience yields have predictive ability beyond that of commodity prices. The macroeconomic variable that I will control for as exchange rate predictors are the standard fundamentals of the sticky-price monetary model (see Cheung et al., 2005): differential of money supply growth, industrial production growth, inflation and interest rate relative to the base country.

The fact that commodities are priced in U.S. dollars has the potential to create an endogeneity problem when U.S. dollar based exchange rates are used for e_t. The intuition is the following. Suppose that the U.S. dollar is hit by a negative (depreciation) shock, independent of developments on commodity markets. This causes the price of a commodity converted in other currencies to decrease one to one. To accommodate this shock the equilibrium quantity and/or price of the commodity increase. The higher demand for the commodity then has a positive impact on its convenience yield, as the probability of an aggregate shortage increases. As a result, commodity convenience yields are positively correlated with exchange rates based on the U.S. dollar. As the coefficient $\beta^{(h)}$ in (8) is expected to have a negative sign, using U.S. dollar based exchange rates would bias the estimate towards zero. I therefore depart from the standard practice in the empirical exchange rate literature and consider the British pound as the base currency instead. In a robustness exercise, I will show the bias on the estimated coefficient when U.S. dollar based exchange rates are used.

A useful empirical characteristic of commodity convenience yields for forecasting commodity currencies is that they are persistent but stationary variables, whereas commodity prices are generally non-stationary[1]. Engel et al. (2010) show that in present-value asset pricing models with discount factors close to unity, asset prices behave like a random walk and short-horizon regressions on fundamentals display low R^2. This is the observed behavior of exchange rates and macroeconomic fundamentals. However in the presence of stationary fundamentals, long-horizon regression can have substantial power, even when the discount factor is close to one and the power of short-horizon is low. In the framework of Engel et al. (2010), the long run level of an asset price is determined by the non-stationary fundamentals. However, the asset price can substantially deviate from its long-run level and revert only gradually due to the stationary fundamentals. This explains why a stationary fundamental, being the short-run deviation of the asset price from its long-run equilibrium

[1] I provide evidence on this in Section 4.1

level, can forecast the movement of the asset price at medium- to long-horizons.

Compared to models explaining one period ahead changes in the exchange rate, inference in predictive regressions with longer horizons ($h > 1$) comes with additional difficulties. The fact that successive observations of the dependent variable are overlapping generates strong serial correlation in the error term $\epsilon_{t+h}^{(h)}$. Standard errors that do not account for this fact will lead to biased inference. A typical solution used in the literature to address this issue has been to use autocorrelation robust estimates of the standard errors such as those proposed in Newey and West (1987). A drawback of this procedure is that it tends to perform poorly in finite samples, leading to rejection rates of the null of no predictability above the nominal level, because it does not capture all of the serial correlation induced by overlapping observations.

I will therefore complement my baseline analysis using Newey-West HAC standard errors with an additional inference techniques. A simple method proposed recently by Hjalmarsson (2011), which is found to have good small sample properties, consists in dividing the standard $t-$statistic by the square root of the forecasting horizon to correct for the effect of overlap in the dependent variable (see Appendix A). This scaled $t-$statistic is then compared to the usual critical value from a $t-$distribution to test the hypothesis of no predictability.

3.3 Data

I consider three commodity currencies relative to the British pound to evaluate the predictive ability of commodity convenience yields for exchange rates: the Australian, Canadian and New Zealand dollars. Among commodity exporters, these three countries have the longest experience with floating exchange rates and each of them exports a variety of goods, which makes their exchange rate responsive to different shocks. I also consider the U.S. dollar as a base currency in a robustness exercise to evaluate whether the fact that commodities are priced in U.S. dollar leads to a bias in the estimated coefficient as argued above. I use end-of-month exchange rates from IMF's International Financial Statistics over a sample period that starts at a different date for each currency but ends in January 2012 for all three countries. For each currency the first observation is the earliest of two dates: the first month after the currency was floated or the first month for which crude oil futures prices are available (April 1983). This corresponds to April 1983 for Canada, January 1984 for Australia and April 1985 for New Zealand. The rationale behind this sample selection is to have the longest possible sample, hence reducing the small sample bias that prevails in fundamentals-based exchange rates models (see Bacchetta et al., 2010) while taking into account the central role of oil prices in commodity markets.

The series of commodity futures prices are created from historical price data on successive futures contracts obtained from Norgate Investor Services and are at a daily frequency from April 1983 to January 2012. Prices at the monthly frequency are calculated as the average of daily prices. The data set contains 20 commodities of different types such as agricultural products, metals or energy and are detailed in the Appendix. Data availability constrains the choice of the commodities included as not every commodity has a long enough history of future contracts. As a consequence some commodities exported by Australia, Canada or New Zealand are not included in the dataset. This is not a problem if the commodities excluded are affected by the same shocks than the commodities included in the panel.

The measure of convenience yield used in this paper differs from the theoretical measure provided by (3) on two dimensions. Firstly, the spot price of a commodity P_t is approximated by the price of the nearest futures contract that is traded $F_{1,t}$, as spot markets often lack the necessary liquidity to provide the correct price for immediate delivery of a commodity. Consequently, the futures price F_t is approximated by the price of the second nearest futures contract $F_{2,t}$. An advantage of this procedure is that the two prices pertain to the same specification of the commodity, e.g. in terms of quality, quantity and delivery conditions. This is not necessarily the case if one compares a spot price and a price of a future contract. Secondly, as there are not necessarily futures contracts of a given commodity expiring every month, the time separating the maturity of the nearest and the second nearest futures contracts varies over the year[2]. For example, futures contracts for corn traded on the Chicago Board of Trade (CBOT) have five delivery months every year: March (H), May (K), July (N), September (U) and December (Z). In January the spot price is approximated by the price of the March contract, whereas the futures price is approximated by the price of the May contract and two months separate the two contracts. In August, the spot price is the price of the September contract and the futures price is the price of the December contract, with three months separating the two contracts.

The convenience yield for each commodity is approximated by the net percentage convenience yield (following Gospodinov and Ng (2011)) and computed as

$$\varphi_t^j = \frac{(1+i_t)F_{1,t}^j - F_{2,t}^j}{F_{1,t}^j} * \frac{1}{G_t^j} \tag{9}$$

where i_t is the return of a three-month U.S. Treasury bill adjusted for the time separating the nearest and second nearest futures contract. The first term corresponds to the convenience yield earned over the whole period separating the nearest and second nearest futures contract. It is divided by the time separating the two contracts G_t^j to obtain a convenience yield corresponding to a period of one month.

The fact that futures contracts do not mature every month might induce errors in the measurement of convenience yields at the monthly frequency. A classical attenuation bias arises with $\beta^{(h)}$ biased towards zero if these errors are uncorrelated with the dependent variable. To assess the extend of the measurement error, I use two commodities - crude oil and heating oil - that have futures contracts maturing every month. I compute the correlation between the convenience yield calculated using the full set of futures contracts and the convenience yield calculated using a restricted set of futures contracts mimicking the contracts available for other commodities, e.g. March (H), May (K), July (N), September (U) and December (Z) in the case of corn. Depending on the set of futures contracts used,

[2]Table A1 in the Appendix describes the contract specification for each commodity.

the correlation between the original and the counterfactual convenience yield ranges from 0.905 to 0.972 for crude oil and from 0.623 to 0.927 for heating oil. The high level of correlation indicates that the measure of convenience yield can be considered as a good proxy even for commodities that do not have futures contracts maturing every month.

The benchmark model against which the predictive ability of convenience yields is assessed uses commodity prices as a predictor. These are country-specific export-weighted indices of commodity prices obtained from the Reserve Bank of Australia, the Bank of Canada and ANZ Bank for New Zealand. The macroeconomic fundamentals are from International Financial Statistics of the IMF and OECD's Main Economic Indicators and are described in detail in Appendix B.2.

4 Empirical results

In this section, I investigate empirically the predictive ability of commodity convenience yields for the exchange rate of commodity currencies. I first provide some descriptive statistics on commodity prices and convenience yields and motivate the predictive ability of convenience yields for exchange rates by providing evidence that commodity convenience yields have predictive ability for commodity returns. Then I present the results of the regression analysis. In the last section I evaluate the out-of-sample forecast performance of the convenience yields model.

4.1 Commodity prices and convenience yields

The picture that emerges from the descriptive statistics is that of highly persistent (and possibly non-stationary) spot commodity prices with first-order autocorrelation coefficients above .97 for all commodities except for *cotton, hogs* and *lumber* and with a maximum value of .997 (*gold*). Accordingly, first-differences of commodity prices display low autocorrelation. Convenience yields are also persistent with first-order autocorrelation coefficients between

Table 1. Descriptive statistics

Commodity	mean(ΔS_t^j)	corr($\Delta S_t^j, \Delta S_{t-1}^j$)	corr(S_t^j, S_{t-1}^j)	mean(φ_t^j)	corr($\varphi_t^j, \varphi_{t-1}^j$)
Agricultural					
Cocoa	0.001	0.212	0.985	-0.004	0.909
Coffee	0.002	0.202	0.979	-0.005	0.948
Corn	0.002	0.326	0.979	-0.011	0.845
Cotton	0.001	0.367	0.963	-0.001	0.751
Feeder cattle	0.002	0.285	0.987	0.007	0.735
Hogs	0.001	0.239	0.908	-0.011	0.666
Live cattle	0.002	0.255	0.973	0.008	0.723
Lumber	0.001	0.146	0.964	-0.012	0.797
Oats	0.002	0.131	0.974	-0.011	0.832
Orange juice	0.002	0.236	0.972	0.001	0.885
Soybean	0.002	0.254	0.976	0.002	0.588
Soybean oil	0.003	0.320	0.982	-0.001	0.928
Sugar	0.004	0.357	0.978	0.003	0.889
Wheat	0.002	0.198	0.980	-0.006	0.868
Metals					
Gold	0.004	0.095	0.997	-0.003	0.520
Palladium	0.005	0.246	0.992	0.011	0.557
Platinum	0.004	0.247	0.995	0.010	0.626
Silver	0.003	0.127	0.993	-0.006	0.910
Energy					
Crude oil	0.003	0.295	0.990	0.006	0.862
Heating oil	0.004	0.268	0.991	0.005	0.798

Notes: The Table presents descriptive statistics for commodity spot prices S_t^j and convenience yields φ_t^j. The sample period is April 1983 to January 2012.

.52 (*gold*) and 0.948 (*coffee*).

As outlined in the previous section, a central argument for the predictive ability of convenience yields for commodity currencies exchange rates is that convenience yields have predictive ability for commodity price changes. Table 2 provides evidence on the predictive ability of individual commodity convenience yields for commodity returns at different horizons. The empirical model estimated by OLS is

$$cp_{t+h}^j - cp_t^j = c^{(h)} + \alpha^{(h)}(L)\Delta cp_t^j + \beta^{(h)}\varphi_t^j + \gamma^{(h)} i_t^{(h)} + \epsilon_{t+h}^{(h)} \tag{10}$$

where cp_t^j is the log price of commodity j, φ^j is the convenience yields of commodity j and $i_t^{(h)}$ the nominal interest rate over period h. The t-statistics computed using Newey-West

HAC standard errors estimates with 24 lags are shown in parentheses below the estimates of $\beta^{(h)}$.

The results indicate that individual price changes of agricultural and energy commodities are negatively and significantly related to convenience yields. A high convenience yields predicts a decrease in the price of the commodity. Overall, the evidence of predictability is strongest for short to medium horizons, although convenience yields have predictive ability for returns at a horizon of 24 months for about half of the commodities. In contrast, the returns on metals prices do not seem to be significantly related to the level of convenience yields.

The first two principal components of commodity convenience yields (denoted by φ^{pc1} and φ^{pc2} below) explain about 30% of the variance in the panel of individial convenience yields, with 19% alone for the first component. The weights used to calculate the principal component scores from individual convenience yields (standardized to have a mean of 0 and standard deviation of 1) are provided in Table A2 in the Appendix. The first principal component loads homogeneously on the individual commodities and is approximately an unweighted average of the individual convenience yields. The second principal component puts large weights on metals and energy and negative weights on most agricultural commodities. φ^{pc1} and φ^{pc2} inherit the persistence of individual convenience yield and display autocorrelation coefficients of .94 and .89, respectively.

4.2 The predictive ability of convenience yields for commodity currencies

I report the estimation of the empirical exchange rate equation (8) for the Australian, Canadian and New Zealand dollars versus the British pound in Tables 2 to 4. Each column presents the OLS estimates of (8) with the dependent variable, the change in the (log-)

Table 2. The predictive power of convenience yields for commodity returns

Dependent variable = $\Delta^h cp^j_{t+h}$	Horizon (in months)					
	(1)	(3)	(6)	(12)	(18)	(24)
Agricultural						
Cocoa	-0.328	-0.730	-0.549	-1.841	-2.934+	-5.501**
	(-1.505)	(-1.018)	(-0.493)	(-1.238)	(-1.900)	(-2.644)
Coffee	-0.065	-0.097	0.006	-0.428	-0.723	-1.596
	(-0.496)	(-0.238)	(0.007)	(-0.293)	(-0.369)	(-0.754)
Corn	-0.581**	-1.789**	-2.793**	-3.107**	-4.423**	-4.544**
	(-6.232)	(-7.572)	(-7.185)	(-5.565)	(-5.601)	(-4.192)
Cotton	-0.478**	-1.188**	-0.824*	-0.241	-0.707	-0.731
	(-6.239)	(-6.603)	(-2.242)	(-0.528)	(-1.126)	(-1.292)
Feeder Cattle	-0.522**	-1.303**	-1.017+	-0.192	-1.186	-1.445
	(-3.722)	(-2.892)	(-1.744)	(-0.240)	(-1.180)	(-1.365)
Hogs	-0.540**	-1.167**	-0.723**	-0.734**	-1.315**	-1.168**
	(-10.614)	(-11.037)	(-4.351)	(-3.551)	(-4.959)	(-4.995)
Live Cattle	-0.540**	-0.897**	-0.154	-0.021	-0.047	-0.203
	(-8.512)	(-5.296)	(-0.745)	(-0.071)	(-0.121)	(-0.523)
Lumber	-0.408**	-0.896**	-0.427	-0.891	-1.083	-1.396*
	(-4.200)	(-3.485)	(-1.351)	(-1.524)	(-1.470)	(-2.183)
Oats	-0.303**	-0.672*	-0.732+	-1.502**	-2.484**	-2.476**
	(-3.394)	(-2.588)	(-1.717)	(-2.856)	(-3.880)	(-3.594)
Orange Juice	-0.426**	-1.385**	-2.305**	-3.804**	-4.800**	-5.567**
	(-3.486)	(-3.399)	(-2.980)	(-3.431)	(-3.833)	(-3.546)
Soybeans	-0.796**	-1.935**	-2.106**	-2.011**	-2.655**	-2.618**
	(-11.694)	(-8.932)	(-4.871)	(-3.668)	(-4.562)	(-3.480)
Soybean Oil	-0.532**	-1.855**	-3.473**	-6.402**	-10.439**	-12.479**
	(-2.712)	(-3.150)	(-3.526)	(-4.450)	(-5.131)	(-6.071)
Sugar	-0.440**	-1.300**	-1.686**	-2.346**	-3.475**	-3.470**
	(-5.769)	(-5.086)	(-3.512)	(-3.529)	(-4.508)	(-3.691)
Wheat	-0.283**	-0.904**	-1.250**	-1.500*	-2.366*	-1.960
	(-4.007)	(-4.653)	(-3.439)	(-2.204)	(-2.248)	(-1.481)
Metals						
Gold	-1.016	-2.390	-1.760	-3.599	-7.965	-14.215
	(-1.066)	(-1.464)	(-0.512)	(-0.533)	(-0.677)	(-1.030)
Palladium	-0.898	-0.952	-1.184	0.359	1.689	3.418
	(-1.564)	(-1.211)	(-0.934)	(0.196)	(0.616)	(0.887)
Platinum	0.247	1.180	1.792	1.287	0.156	-0.473
	(0.681)	(1.280)	(1.198)	(0.611)	(0.067)	(-0.185)
Silver	-2.196	-4.051	-3.284	-2.043	-6.021	-4.859
	(-1.354)	(-1.343)	(-0.874)	(-0.412)	(-0.921)	(-0.653)
Energy						
Crude Oil	-0.766**	-2.155*	-2.368+	-4.456**	-3.317	-3.829
	(-2.659)	(-2.347)	(-1.811)	(-2.824)	(-1.519)	(-1.456)
Heating Oil	-0.718**	-1.460**	-0.838	-2.580**	-2.870*	-3.854**
	(-4.396)	(-2.953)	(-1.379)	(-3.312)	(-2.457)	(-2.634)

Notes: The Table presents OLS estimates of the coefficient $\beta^{(h)}$ in the regression $cp^j_{t+h} - cp^j_t = c^{(h)} + \alpha^{(h)}(L)\Delta cp^j_t + \beta^{(h)}\varphi^j_t + \gamma^{(h)} i^{(h)}_t + \epsilon^{(h)}_{t+h}$, where cp^j_t is the log price of commodity j, φ^j is the convenience yields of commodity j and $i^{(h)}_t$ the nominal interest rate over period h. All regressions include two additional lags of the first-differences of cp^j_t and a constant (coefficients estimates not reported). The t-statistic computed using Newey-West HAC standard errors estimates with 24 lags are shown in parentheses below the coefficients estimates. Levels of significance indicated by ** p<0.01, * p<0.05, + p<0.1

exchange rate, calculated over a different horizon. The horizons considered are 1, 3, 6, 12, 18 and 24 months. All regressions include the variables of interest, i.e. the first two principal components of commodity convenience yields φ_t^{pc1} and φ_t^{pc2}, three lags of the first-differences of the exchange rate Δe_t and of the country-specific index of commodity prices Δcp_t and a constant[3]. For each horizon, the regression in the second column also includes macroeconomic fundamentals: differential of money supply growth, industrial production growth, inflation and interest rate relative to the base country. t-statistics computed using Newey-West HAC standard errors estimates with 24 lags are shown in parentheses below the coefficients estimates.

We first observe that the coefficient estimate on the variable of interest has a negative sign, which is consistent with the evidence that i) commodity prices and commodity currencies' exchange rate are positively related and ii) a high convenience yield predicts a decrease in commodity prices. The value of the coefficient increases (in absolute terms) with the forecast horizon and so does the R^2 of the regression. The coefficient on the first principal component of convenience yields is significant at the 1% level for the Australian dollar at all horizons, at the 1% level for the Canadian dollar at horizons of 12 months or more and at the 5% level for the New Zealand dollar at all horizons.

As mentioned in the section on the empirical methodology, inference based on Newey-West standard errors estimates in long-horizon regressions with overlapping observations has been criticized on its small sample properties. To complement my investigation I calculate scaled t-statistics following Hjalmarsson (2011) which are given in the second line of numbers in parentheses under the coefficient estimates of φ_t^{pc1}. A first finding is that the scaled t-statistics are lower than the t-statistics calculated using Newey-West standard errors leading to higher significance level and fewer rejections of the null hypothesis of no predictability. We observe that convenience yields keep their predictive power for the Australian dollar at a significance level below 1%. The evidence for the Canadian and New Zealand dollars are

[3]The coefficient estimates of the second principal component (not significant), the two last lags of Δe_t and of Δcp_t are not reported to save space

more mixed with significance levels between 1% and 10%. Together, these findings indicate that a high level of convenience yields, which might be the consequence of low commodity inventory levels and/or high uncertainty about future demand, predicts a depreciation of the Australian, Canadian and New Zealand dollar against the British pound.

The estimated coefficients on the lagged commodity price index changes are significant and positive only in a few cases. This result is in line with Chen et al. (2010) who find only weak evidence of Granger-causality from commodity price indices to exchange rates and Ferraro et al. (2011) who find little systematic relation between oil prices and the exchange rate at the monthly frequency. These results show the importance of using variables that have predictive power for commodity returns, such as convenience yields to predict exchange rate returns. In line with the literature, the monetary fundamentals are insignificant in most cases.

Table 3. The predictive ability of convenience yields for the Australian Dollar

Dependent variable = $\Delta^h e_{t+h}$	Horizon (in months)											
	(1)	(1)	(3)	(3)	(6)	(6)	(12)	(12)	(18)	(18)	(24)	(24)
φ_t^{pc1}	**-0.006**	**-0.008**	**-0.016**	**-0.018**	**-0.027**	**-0.031**	**-0.062**	**-0.063**	**-0.093**	**-0.096**	**-0.109**	**-0.111**
(t^{NW})	**(-3.492)****	**(-3.492)****	**(-2.937)****	**(-2.919)****	**(-3.236)****	**(-3.181)****	**(-4.707)****	**(-3.978)****	**(-5.545)****	**(-5.362)****	**(-5.264)****	**(-5.347)****
(t^{scaled})	**(-3.262)****	**(-3.855)****	**(-2.699)****	**(-2.829)****	**(-2.360)***	**(-2.600)****	**(-3.152)****	**(-3.044)****	**(-3.714)****	**(-3.547)****	**(-3.172)****	**(-2.999)****
Δcp_t^{AUS}	-0.037	-0.037	0.025	0.020	-0.197	-0.193	-0.107	-0.192	0.110	0.003	-0.074	-0.189
	(-0.419)	(-0.406)	(0.225)	(0.177)	(-0.917)	(-0.902)	(-0.289)	(-0.562)	(0.313)	(0.011)	(-0.173)	(-0.489)
Δe_t	0.032	0.034	-0.100	-0.095	-0.184+	-0.178*	-0.262*	-0.295**	-0.333**	-0.328*	-0.315*	-0.315+
	(0.374)	(0.396)	(-1.033)	(-0.986)	(-1.932)	(-2.006)	(-2.582)	(-2.976)	(-2.708)	(-2.592)	(-2.244)	(-1.843)
$\Delta(m_t - m_t^*)$		-0.028+		-0.029+		-0.056*		0.042		0.030		0.060
		(-1.917)		(-1.840)		(-2.296)		(0.942)		(0.520)		(1.009)
$\Delta(p_t - p_t^*)$		0.225		-0.052		0.064		-0.825		-1.944		-1.712
		(0.555)		(-0.054)		(0.067)		(-0.662)		(-1.547)		(-1.253)
$\Delta(y_t - y_t^*)$		0.066		0.257		0.628+		0.049		0.672+		0.837**
		(0.454)		(0.962)		(1.746)		(0.136)		(1.875)		(2.630)
$i_t - i_t^*$		1.678		2.047		3.879		4.568		10.983		12.824
		(1.643)		(0.769)		(0.719)		(0.618)		(1.306)		(1.184)
Observations	336	332	334	330	331	327	325	321	319	315	313	309
R^2	0.039	0.068	0.074	0.087	0.100	0.136	0.283	0.294	0.460	0.460	0.450	0.454

Notes: The Table presents OLS estimates of the regression $e_{t+h} - e_t = c^{(h)} + \alpha^{(h)}(L)\Delta e_t + \beta^{(h)}\varphi_t + \gamma^{(h)}(L)\Delta cp_t + \delta^{(h)}X_t + \epsilon_{t+h}^{(h)}$, where e_t is the log-exchange rate of the Australian dollar versus the British pound, φ^{pc1} is the first principal component of commodity convenience yields and cp_t the country-specific index of commodity prices. X_t contains standard macro fundamentals: $\Delta(m_t - m_t^*)$ is the differential in money growth, $\Delta(p_t - p_t^*)$ the differential in inflation, $\Delta(y_t - y_t^*)$ the differential in industrial production growth and $i_t - i_t^*$ the interest rate differential between Australia and the United Kingdom. All regressions include two additional lags of the first-differences of e_t and cp_t and a constant (coeffificents estimates not reported). The t-statistic computed using Newey-West HAC standard errors estimates with 24 lags are shown in parentheses below the coefficients estimates. The second line of numbers in parentheses under the coefficient estimates of φ^{pc1} is the scaled t-statistic computed following Hjalmarsson (2011). Levels of significance indicated by ** $p<0.01$, * $p<0.05$, + $p<0.1$

TABLE 4. THE PREDICTIVE ABILITY OF CONVENIENCE YIELDS FOR THE CANADIAN DOLLAR

Dependent variable = $\Delta^h e_{t+h}$	Horizon (in months)											
	(1)	(1)	(3)	(3)	(6)	(6)	(12)	(12)	(18)	(18)	(24)	(24)
φ_t^{pc1}	**-0.002**	**-0.002**	**-0.003**	**-0.004**	**-0.007**	**-0.012**	**-0.028**	**-0.035**	**-0.049**	**-0.053**	**-0.063**	**-0.063**
(t^{NW})	**(-1.270)**	**(-1.333)**	**(-0.596)**	**(-0.979)**	**(-0.899)**	**(-1.458)**	**(-2.772)****	**(-2.970)****	**(-4.292)****	**(-4.043)****	**(-4.359)****	**(-4.010)****
(t^{scaled})	**(-1.120)**	**(-1.265)**	**(-0.520)**	**(-0.867)**	**(-0.697)**	**(-1.135)**	**(-1.639)**	**(-1.966)***	**(-2.197)***	**(-2.367)***	**(-2.343)***	**(-2.224)***
Δcp_t^{CAN}	-0.039	-0.042	0.017	0.006	0.087	0.077	0.212*	0.198+	0.178	0.196	0.103	0.094
	(-0.874)	(-0.886)	(0.204)	(0.064)	(0.850)	(0.727)	(1.970)	(1.732)	(0.922)	(1.057)	(0.745)	(0.698)
Δe_t	-0.052	-0.068	-0.072	-0.084	-0.140	-0.170+	-0.242*	-0.271**	-0.490**	-0.462**	-0.314*	-0.245
	(-1.029)	(-1.243)	(-1.016)	(-1.097)	(-1.398)	(-1.922)	(-2.361)	(-2.640)	(-4.826)	(-4.118)	(-2.175)	(-1.582)
$\Delta(m_t - m_t^*)$		-0.015*		-0.032**		-0.069**		-0.015		0.000		0.013
		(-2.144)		(-4.236)		(-5.138)		(-0.508)		(0.002)		(0.331)
$\Delta(p_t - p_t^*)$		0.265		1.373*		0.913		1.050+		-0.445		1.414+
		(1.110)		(2.434)		(1.421)		(1.966)		(-0.757)		(1.687)
$\Delta(y_t - y_t^*)$		0.051		0.053		0.226		0.015		0.451*		0.018
		(0.430)		(0.309)		(0.764)		(0.039)		(2.007)		(0.060)
$i_t - i_t^*$		3.137		1.655		2.603		2.072		-6.105		-8.354
		(1.585)		(0.396)		(0.422)		(0.240)		(-0.661)		(-1.065)
Observations	345	331	343	330	340	327	334	321	328	315	322	309
R^2	0.012	0.037	0.008	0.045	0.030	0.073	0.124	0.175	0.242	0.292	0.308	0.313

Notes: The Table presents OLS estimates of the regression $e_{t+h} - e_t = c^{(h)} + \alpha^{(h)}(L)\Delta e_t + \beta^{(h)}\varphi_t + \gamma^{(h)}(L)\Delta cp_t + \delta^{(h)}X_t + \epsilon_{t+h}^{(h)}$, where e_t is the log-exchange rate of the Canadian dollar versus the British pound, φ^{pc1} is the first principal component of commodity convenience yields and cp_t the country-specific index of commodity prices. X_t contains standard macro fundamentals: $\Delta(m_t - m_t^*)$ is the differential in money growth, $\Delta(p_t - p_t^*)$ the differential in inflation, $\Delta(y_t - y_t^*)$ the differential in industrial production growth and $i_t - i_t^*$ the interest rate differential between Canada and the United Kingdom. All regressions include two additional lags of the first-differences of e_t and cp_t and a constant (coeffificents estimates not reported). The t-statistic computed using Newey-West HAC standard errors estimates with 24 lags are shown in parentheses below the coefficients estimates. The second line of numbers in parentheses under the coefficient estimates of φ^{pc1} is the scaled t-statistic computed following Hjalmarsson (2011). Levels of significance indicated by ** p<0.01, * p<0.05, + p<0.1

Table 5. The predictive ability of convenience yields for the New Zealand Dollar

Dependent variable = $\Delta^h e_{t+h}$	Horizon (in months)											
	(1)	(1)	(3)	(3)	(6)	(6)	(12)	(12)	(18)	(18)	(24)	(24)
φ_t^{pc1}	**-0.004**	**-0.004**	**-0.009**	**-0.008**	**-0.016**	**-0.016**	**-0.039**	**-0.038**	**-0.069**	**-0.068**	**-0.086**	**-0.085**
(t^{NW})	**(-2.812)****	**(-2.523)***	**(-2.459)***	**(-2.066)***	**(-2.871)****	**(-2.369)***	**(-3.177)****	**(-2.874)****	**(-4.025)****	**(-3.927)****	**(-3.775)****	**(-3.832)****
(t^{scaled})	**(-1.939)**+	**(-1.696)**+	**(-1.557)**	**(-1.315)**	**(-1.518)**	**(-1.383)**	**(-1.906)**+	**(-1.775)**+	**(-2.458)***	**(-2.322)***	**(-2.224)***	**(-2.099)***
Δcp_t^{NZ}	0.053	0.038	0.331**	0.314**	0.352**	0.316**	0.102	0.066	0.049	0.003	0.079	0.030
	(0.783)	(0.525)	(3.043)	(2.729)	(3.117)	(2.813)	(0.477)	(0.295)	(0.166)	(0.011)	(0.264)	(0.099)
Δe_t	-0.099	-0.095	-0.167	-0.167	-0.245*	-0.237*	-0.217	-0.212	-0.233	-0.218	-0.177	-0.158
	(-1.582)	(-1.462)	(-1.498)	(-1.464)	(-2.252)	(-2.119)	(-1.524)	(-1.438)	(-1.467)	(-1.360)	(-1.037)	(-0.910)
$\Delta(m_t - m_t*)$		0.049		0.078		0.094		0.172		0.134		-0.279
		(0.850)		(0.647)		(0.564)		(0.769)		(0.571)		(-0.938)
$\Delta(p_t - p_t*)$		0.060		0.643		0.702		0.444		0.548		1.589+
		(0.189)		(0.968)		(1.403)		(0.568)		(0.624)		(1.965)
$\Delta(y_t - y_t*)$		-0.142		-0.100		0.224		0.230		0.790**		0.633+
		(-1.242)		(-0.458)		(0.845)		(0.600)		(2.805)		(1.779)
$i_t - i_t^*$		-0.165		-1.258		-1.033		-0.783		-1.716		-4.102
		(-0.241)		(-0.685)		(-0.408)		(-0.162)		(-0.269)		(-0.521)
Observations	309	306	307	306	304	304	298	298	292	292	286	286
R^2	0.047	0.053	0.073	0.081	0.110	0.115	0.208	0.211	0.337	0.347	0.335	0.346

Notes: The Table presents OLS estimates of the regression $e_{t+h} - e_t = c^{(h)} + \alpha^{(h)}(L)\Delta e_t + \beta^{(h)}\varphi_t + \gamma^{(h)}(L)\Delta cp_t + \delta^{(h)} X_t + \epsilon_{t+h}^{(h)}$, where e_t is the log-exchange rate of the New Zealand dollar versus the British pound, φ^{pc1} is the first principal component of commodity convenience yields and cp_t the country-specific index of commodity prices. X_t contains standard macro fundamentals: $\Delta(m_t - m_t^*)$ is the differential in money growth, $\Delta(p_t - p_t^*)$ the differential in inflation, $\Delta(y_t - y_t^*)$ the differential in industrial production growth and $i_t - i_t^*$ the interest rate differential between New Zealand and the United Kingdom. All regressions include two additional lags of the first-differences of e_t and cp_t and a constant (coeffificents estimates not reported). The t-statistic computed using Newey-West HAC standard errors estimates with 24 lags are shown in parentheses below the coefficients estimates. The second line of numbers in parentheses under the coefficient estimates of φ^{pc1} is the scaled t-statistic computed following Hjalmarsson (2011). Levels of significance indicated by ** p<0.01, * p<0.05, + p<0.1

In Table 6, I report the estimation of equation (8) when country-specific convenience yield indices are used as predictors instead of the principal component of convenience yields[4]. The use of a different aggregate measure of convenience yields does not have a qualitative impact as the estimated coefficient is still negative, except in a few cases. Unsurprisingly, as the country-specific indices do not cover the whole spectrum of commodities exported by each country, the significance of the results is strongly affected, especially for the Canadian and New Zealand dollar.

4.2.1 U.S. dollar as base currency

The preceding section shows evidence that commodity convenience yields incorporate information useful for the prediction of the British pound based exchange rate of three commodity exporters. In this section I explore whether the observed relationship between movements in the value of the commodity currencies and convenience yields are affected by the value of the U.S. dollar, as both the currencies and commodities are priced in this currency. To address this issue I repeat the analysis of the preceding section and estimate the exchange rate equation (8) with exchange rates based on the U.S. dollar as a dependent variable and including all control variables.

The results reported in Table 7 confirm the ability of commodity convenience yields to predict changes in the exchange rates of commodity currencies. The estimated coefficient on the variable of interest is negative and highly significantly so for the Australian dollar. However, the significance of the coefficient is weaker overall, especially when considering the more conservative scaled t-statistic for the Canadian and New Zealand dollars. These results are in line with the claim made in Section 3.2 that an attenuation bias might affect the estimated coefficient on the convenience yield when the exchange rates are based on the U.S. dollar.

[4]Regressions involving the simple average of commodity convenience yields instead of the principal components yield qualitatively similar results. The results are available upon request.

Table 6. The predictive ability of country-specific convenience yields indices

Dependent variable = $\Delta^h e_{t+h}$	Horizon (in months)					
	(1)	(3)	(6)	(12)	(18)	(24)
	A. Australian dollar vs. British pound					
φ_t^{AUS}	**-0.525**	**-1.456**	**-2.516**	**-3.738**	**-5.891**	**-7.287**
(t^{NW})	**(-1.930)**+	**(-1.857)**+	**(-2.213)***	**(-1.609)**	**(-1.887)**+	**(-1.938)**+
(t^{scaled})	**(-1.842)**+	**(-1.679)**+	**(-1.509)**	**(-1.171)**	**(-1.290)**	**(-1.186)**
Δcp_t^{AUS}	-0.028	0.042	-0.148	-0.063	0.175	-0.053
	(-0.300)	(0.357)	(-0.531)	(-0.118)	(0.306)	(-0.083)
Δe_t	0.054	-0.061	-0.115	-0.150	-0.122	-0.090
	(0.583)	(-0.568)	(-1.083)	(-1.073)	(-0.615)	(-0.430)
Observations	332	330	327	321	315	309
R^2	0.031	0.044	0.065	0.088	0.139	0.150
	B. Canadian dollar vs. British pound					
φ_t^{CAN}	**-0.040**	**-0.205**	**-0.489**	**-2.034**	**-2.754**	**-3.394**
	(-0.396)	**(-0.725)**	**(-1.033)**	**(-3.242)****	**(-3.418)****	**(-2.604)****
	(-0.292)	**(-0.512)**	**(-0.605)**	**(-1.430)**	**(-1.428)**	**(-1.369)**
Δcp_t^{CAN}	-0.040	0.009	0.085	0.207	0.200	0.095
	(-0.827)	(0.100)	(0.739)	(1.395)	(0.843)	(0.514)
Δs_t	-0.062	-0.073	-0.147	-0.245*	-0.428**	-0.221
	(-1.148)	(-0.936)	(-1.614)	(-2.334)	(-3.463)	(-1.229)
Observations	331	330	327	321	315	309
R^2	0.028	0.037	0.050	0.120	0.159	0.162
	C. New Zealand dollar vs. British pound					
φ_t^{pc1}	**-0.033**	**0.069**	**-0.230**	**0.100**	**-1.335**	**-0.895**
	(-0.217)	**(0.175)**	**-0.324)**	**(0.077)**	**(-0.775)**	**(-0.383)**
	(-0.173)	**(0.135)**	**(-0.231)**	**(0.048)**	**(-0.432)**	**(-0.211)**
Δcp_t^{NZ}	0.040	0.309**	0.313*	0.069	0.022	0.082
	(0.555)	(2.662)	(2.405)	(0.235)	(0.057)	(0.197)
Δs_t	-0.090	-0.155	-0.203	-0.155	-0.129	-0.088
	(-1.407)	(-1.320)	(-1.611)	(-0.981)	(-0.705)	(-0.406)
Observations	306	306	304	298	292	286
R^2	0.039	0.044	0.034	0.018	0.037	0.025

Notes: The Table presents OLS estimates of the regression $e_{t+h} - e_t = c^{(h)} + \alpha^{(h)}(L)\Delta e_t + \beta^{(h)}\varphi_t + \gamma^{(h)}(L)\Delta cp_t + \epsilon_{t+h}^{(h)}$, where e_t is the log-exchange rate of either the Australian, Canadian or New Zealand dollar versus the British pound, φ^j is the country-specific convenience yield index of country j = AUS, CAN, NZ and cp_t the country-specific index of commodity prices. All regressions include two additional lags of the first-differences of e_t and cp_t, standard macroeconomic fundamentals and a constant (coeffificents estimates not reported). The t-statistic computed using Newey-West HAC standard errors estimates with 24 lags are shown in parentheses below the coefficients estimates. The second line of numbers in parentheses under the coefficient estimates of φ^j is the scaled t-statistic computed following Hjalmarsson (2011). Levels of significance indicated by ** p<0.01, * p<0.05, + p<0.1

Table 7. The U.S. dollar as a base currency

Dependent variable = $\Delta^h e_{t+h}$	Horizon (in months) (1)	(3)	(6)	(12)	(18)	(24)
	A. Australian dollar vs. U.S. dollar					
φ_t^{pc1}	**-0.007**	**-0.020**	**-0.034**	**-0.063**	**-0.087**	**-0.097**
(t^{NW})	**(-3.983)****	**(-3.346)****	**(-3.586)****	**(-5.111)****	**(-6.491)****	**(-6.587)****
(t^{scaled})	**(-3.643)****	**(-3.125)****	**(-2.661)****	**(-2.609)****	**(-2.758)****	**(-2.315)****
Δcp_t^{AUS}	-0.194	-0.224+	-0.609+	-0.663	-0.363	-0.485
	(-1.411)	(-1.730)	(-1.809)	(-1.240)	(-1.146)	(-1.023)
Δe_t	0.086	0.150	0.080	-0.119	-0.207	-0.221
	(0.886)	(1.039)	(0.458)	(-0.739)	(-1.212)	(-0.961)
Observations	336	334	331	325	319	313
R^2	0.070	0.118	0.185	0.255	0.367	0.380
	B. Canadian dollar vs. U.S. dollar					
φ_t^{pc1}	**-0.002**	**-0.003**	**-0.006**	**-0.019**	**-0.029**	**-0.034**
	(-2.495)*	**(-1.404)**	**(-1.188)**	**(-2.081)***	**(-2.484)***	**(-2.251)***
	(-1.463)	**(-0.989)**	**(-0.838)**	**(-1.333)**	**(-1.421)**	**(-1.226)**
Δcp_t^{CAN}	0.061*	0.117*	0.073	-0.068	0.025	-0.094
	(2.171)	(2.024)	(1.150)	(-0.455)	(0.258)	(-0.711)
Δs_t	-0.071	-0.027	-0.007	0.209	-0.101	0.242
	(-0.953)	(-0.332)	(-0.050)	(1.263)	(-0.304)	(1.072)
Observations	344	343	340	334	328	322
R^2	0.051	0.051	0.036	0.107	0.143	0.157
	C. New Zealand dollar vs. U.S. dollar					
φ_t^{pc1}	**-0.004**	**-0.009**	**-0.014**	**-0.035**	**-0.058**	**-0.075**
	(-1.818)+	**(-1.684)+**	**(-1.462)**	**(-2.060)***	**(-2.618)****	**(-2.697)****
	(-1.609)	**(-1.342)**	**(-0.957)**	**(-1.123)**	**(-1.275)**	**(-1.263)**
Δcp_t^{NZ}	0.013	0.254	0.101	-0.176	-0.278	-0.393
	(0.138)	(1.617)	(0.531)	(-0.687)	(-0.618)	(-0.770)
Δs_t	-0.014	0.152	0.137	0.092	0.078	0.101
	(-0.203)	(0.890)	(0.548)	(0.447)	(0.275)	(0.289)
Observations	306	306	304	298	292	286
R^2	0.072	0.109	0.141	0.159	0.209	0.227

Notes: The Table presents OLS estimates of the regression $e_{t+h} - e_t = c^{(h)} + \alpha^{(h)}(L)\Delta e_t + \beta^{(h)}\varphi_t + \gamma^{(h)}(L)\Delta cp_t + \epsilon_{t+h}^{(h)}$, where e_t is the log-exchange rate of either the Australian, Canadian or New Zealand dollar versus the U.S. dollar, φ^{pc1} is the first principal component of commodity convenience yields and cp_t the country-specific index of commodity prices. All regressions include two additional lags of the first-differences of e_t and cp_t and a constant (coeffificents estimates not reported). The t-statistic computed using Newey-West HAC standard errors estimates with 24 lags are shown in parentheses below the coefficients estimates. The second line of numbers in parentheses under the coefficient estimates of φ^{pc1} is the scaled t-statistic computed following Hjalmarsson (2011). Levels of significance indicated by ** p<0.01, * p<0.05, + p<0.1

4.3 Out-of-sample performance

Following the tradition of Meese and Rogoff (1983), exchange rate models are evaluated on the basis of their ability to generate accurate out-of-sample predictions. In this section, I evaluate the out-of-sample predictive performance of convenience yields for commodity currencies' exchange rate changes.

I consider a rolling regression scheme to generate out-of-sample exchange rate point forecasts based on the following model

$$\Delta^h e_{t+h} = \alpha^{(h)} + \beta^{(h)} f_t + \epsilon_{t+h}^{(h)} \tag{11}$$

where f_t is a vector of fundamentals.

Forecasts are evaluated on two different periods of length P: either the last 5 years of the sample (P = 60) or the last 10 years (P = 120). In the case of P=60, the first regression is run on a sample of fixed length $L_j - h$, j = AUS, CAN, NZ that ends in January 2007 and the first prediction is for the h-month ahead change from February 2007 using the estimated parameters. For the second prediction, the regression sample is shifted one period later and the prediction made for the h-month ahead change from March 2002, and so on.[5]

I consider four different models. The first model (termed CY) uses the first two principal components of convenience yields as predictors, i.e. φ_t^{pc1} and φ_t^{pc2}. The second model (LCP) is based on three lags of the change in the country-specific index of commodity prices Δcp_t^j, Δcp_{t-1}^j and Δcp_{t-2}^j, j = AUS, CAN, NZ. The third model (CCP) uses the contemporaneous one-period change in the country-specific index of commodity prices Δcp_{t+1}^j, j = AUS, CAN, NZ as a predictor. This model has an information advantage over the first two models, as the fitted value calculated in period t is based on the realized change in the commodity price index between t and $t+1$, whereas the predictions of CY and LCP

[5] As the length of the sample period varies between the three currencies considered, the regression sample and prediction sample length cannot both be equal across currencies. I choose to equalize the length of the prediction sample and let the regression sample length vary across currencies.

are based only on information known in period t. The fourth model (MF) is based on the monetary fundamentals $X_t = [\Delta(m_t - m_t^*), \Delta(p_t - p_t^*), \Delta(y_t - y_t^*), i_t - i_t^*]$.

I evaluate each model on the basis of the percentage difference between the root mean squared errors (RMSE) generated by the model over the simulated out-of-sample sub-period and the RMSE of the benchmark random walk without drift model calculated over the same sub-period. Formally, I calculate the following statistic

$$RMSE^{\%diff} = \frac{RMSE^{\text{model}} - RMSE^{\text{rw}}}{RMSE^{\text{rw}}} \tag{12}$$

with

$$RMSE^{\text{model}} = \sqrt{\frac{1}{P}\sum_{k=0}^{P-1}(\hat{\epsilon}_{t+h+k}^{(h)})^2} \tag{13}$$

$$RMSE^{\text{rw}} = \sqrt{\frac{1}{P}\sum_{k=0}^{P-1}(\Delta^h e_{t+h+k})^2} \tag{14}$$

where $\hat{\epsilon}_{t+h+k}^{(h)} = \Delta^h e_{t+h+k} - \hat{\alpha}^{(h)} - \hat{\beta}^{(h)} f_t$, $\hat{\alpha}^{(h)}$ and $\hat{\beta}^{(h)}$ are the estimated OLS coefficients from the sample including observations $t - L + k$ to $t + k$.

The calculated $RMSE^{\%diff}$ statistics for the four models, three currencies and six different horizons are shown in Table 8. A negative number indicates that the model outperforms the random walk without drift model. The results indicate that the model based on convenience yields consistently and strongly outperforms the three other models and the random walk, except for the Canadian dollar at horizons below 12 months. For one month ahead predictions, the gain in predictive accuracy is highest for the Australian dollar for which it reaches 3% as $RMSE^{\text{CY}}$ is 3% lower than $RMSE^{\text{rw}}$. for the and even 7% for three months ahead forecasts. The gains in predictive accuracy are highest at medium- to long-horizons. When the forecasts are evaluated on the last 10 years of data, the model based on commodity convenience yields is 23% (12 months) to 37% (24 months) more accurate than the random walk for the Australian dollar, 14% to 23% for the Canadian dollar and 9% to 21% for the New Zealand dollar.

The model based on contemporaneous commodity price changes does not consistently forecast exchange rate movements better than a random walk. At short horizons, the CCP model delivers small gains in predictive accuracy for the New Zealand dollar. The largest gains are found at medium to long horizons for the Australian and Canadian dollars. However, unreported results indicate that the CCP model beats the random walk model at horizons of 1 to 6 months when exchange rates are based on the U.S. dollar in line with the findings of Bacchetta et al. (2010). Consistent with Chen et al. (2010) I find that the predictions of the model based on lagged changes in commodity prices are clearly worse than those of the random walk.

5 Conclusion

This paper proposes and evaluates a novel approach to forecast the exchange rate of commodity currencies using commodity convenience yields, which are defined as the yields that accrue to the holder of commodity inventories. The predictive power of commodity convenience yields for the exchange rate of commodity currencies is a consequence of the strong relationship between commodity prices and commodity currencies on the one hand and the forward-looking nature of convenience yields on the other hand. The empirical evidence shown in the paper confirms the role of commodity currencies, the Australian, Canadian and New Zealand dollars, as an exception to the well-documented Meese-Rogoff puzzle.

In-sample, I find that future changes in the bilateral exchange rates of the Australian, the Canadian and New Zealand Dollars vis a vis the British pound or the U.S. dollar are significantly related to aggregate measures of commodity convenience yields. A high level of convenience yields predicts a depreciation of all three exchange rates in a horizon of 12 to 24 months and also at shorter horizon for the Australian and New Zealand dollars.

Table 8. Out-of-sample predictive ability

Predicted variable = $\Delta^h e_{t+h}$	Horizon (in months) (1)	(3)	(6)	(12)	(18)	(24)
			P=60			
Australian dollar vs. Bristish pound						
CY	**−0.031**	**−0.065**	**−0.150**	**−0.393**	**−0.522**	**−0.510**
LCP	0.008	−0.016	−0.021	0.008	0.005	−0.022
CCP	0.040	0.026	0.012	0.007	−0.010	−0.026
MF	0.003	0.004	−0.019	−0.018	−0.056	−0.019
Canadian dollar vs. Bristish pound						
CY	**0.004**	0.034	**−0.002**	**−0.203**	**−0.284**	**−0.317**
LCP	0.021	0.043	0.020	0.060	0.073	−0.035
CCP	0.009	0.033	0.031	0.026	0.022	−0.039
MF	0.016	**0.026**	0.015	−0.012	−0.061	−0.051
New Zealand dollar vs. Bristish pound						
CY	**−0.001**	**−0.049**	**−0.188**	**−0.383**	**−0.458**	**−0.453**
LCP	0.039	0.032	−0.009	0.065	0.033	−0.012
CCP	0.004	0.009	−0.032	−0.001	−0.017	−0.035
MF	0.003	0.000	−0.021	−0.041	−0.003	0.027
			P=120			
Australian dollar vs. Bristish pound						
CY	**−0.018**	**−0.056**	**−0.095**	**−0.233**	**−0.325**	**−0.371**
LCP	0.014	0.001	−0.017	−0.016	−0.008	0.013
CCP	0.026	0.004	0.004	−0.017	−0.020	−0.012
MF	0.008	0.045	−0.008	−0.059	−0.075	−0.056
Canadian dollar vs. Bristish pound						
CY	0.009	0.014	**−0.017**	**−0.136**	**−0.169**	**−0.225**
LCP	0.014	0.017	0.001	0.000	−0.002	−0.022
CCP	**0.004**	0.004	0.003	−0.016	−0.023	−0.043
MF	0.013	**−0.007**	−0.008	0.000	−0.056	−0.055
New Zealand dollar vs. Bristish pound						
CY	**−0.006**	**−0.019**	**−0.031**	**−0.091**	**−0.160**	**−0.212**
LCP	0.030	0.003	0.012	0.045	0.067	0.091
CCP	−0.001	−0.006	−0.018	0.015	0.028	0.048
MF	0.000	0.009	−0.016	−0.026	0.019	0.057

Notes: The table presents the percentage difference between the root mean squared errors (RMSE) generated by each of the models (CY, LCP, CCP and MF) over the simulated out-of-sample sub-period represented by the last P observations of the sample and the RMSE of the random walk without drift model, i.e. $(RMSE^{model} - RMSE^{rw})/(RMSE^{rw})$ using a rolling regression scheme. The four different models are CY : $\Delta^h e^j_{t+h} = \alpha^{(h)} + \beta^{(h)}_1 \varphi^{pc1}_t + \beta^{(h)}_2 \varphi^{pc2}_t + \epsilon^{(h)}_{t+h}$, LCP : $\Delta^h e^j_{t+h} = \alpha^{(h)} + \beta^{(h)}_1 \Delta cp^j_t + \beta^{(h)}_2 \Delta cp^j_{t-1} + \beta^{(h)}_3 \Delta cp^j_{t-2} + \epsilon^{(h)}_{t+h}$, CCP : $\Delta^h e^j_{t+h} = \alpha^{(h)} + \beta^{(h)} \Delta cp^j_{t+1} + \epsilon^{(h)}_{t+h}$ and MF : $\Delta^h e^j_{t+h} = \alpha^{(h)} + \beta^{(h)} X_t + \epsilon^{(h)}_{t+h}$, j = AUS, CAN, NZ, $X_t = [\Delta(m_t - m^*_t), \Delta(p_t - p^*_t), \Delta(y_t - y^*_t), i_t - i^*_t]$. The best performing model for each horizon/currency pair is indicated in bold.

The findings are confirmed in an out-of-sample evaluation exercise. The model based on convenience yields outperforms the random walk without drift model over all horizons. The gains in predictive accuracy reach 37% for the Australian dollar, 23% for the Canadian dollar and 21% for the New Zealand dollar when the forecasts are evaluated over the last 10 years of data.

While this paper has focused on the Australian, Canadian and New Zealand dollars, it would be desirable to look at other commodity currencies, such as the South African rand and Chilean peso, to confirm or invalidate the findings presented in this paper.

Appendix

A Inference in long-horizon regressions

This section presents the inference method in long-horizon regressions proposed by Hjalmarsson (2011). Let the dependent variable be denoted by Δe_t, which represents the one-period exchange rate return, and φ_t the predictor. The long-hozion regression model is written as

$$\Delta^h e_{t+h} = \alpha_h + \beta_h \varphi_t + u_{t+h} \tag{15}$$

where $\Delta^h e_{t+h} = \sum_{j=1}^{h} \Delta e_{t+j}$ and the long-horizon realized return is regressed onto the one-period return. Let $\hat{\beta}_h$ denote the OLS estimator of β_h.

Equation (15) is a fitted regression, whereas the true data-generating model is specified for the one-period return Δe_{t+1} as follows

$$\begin{aligned} \Delta e_{t+1} &= \alpha + \beta \varphi_t + u_{t+1} \\ \varphi_{t+1} &= \gamma + \rho \varphi_t + v_{t+1} \end{aligned}$$

where $\rho = 1 + c/T, t = 1, ..., T$ and T is the sample size. The local-to-unity parameterization of the autoregressive root of the regressor captures the near unit-root or highly persistent behavior of the predictor variable. The errors u_{t+1} and v_{t+1} are assumed to be covariance stationary and satisfy the assumption stated in Appendix B of Hjalmarsson (2011). I consider the case in which u_{t+1} and v_{t+1} are uncorrelated and hence the predictor variable is exogenous to the dependent variable.

Corollary 1 in Hjalmarsson (2011) states that under the null hypothesis of no predictability, for a fixed h as $T \to \infty$

$$\frac{t_h}{\sqrt{h}} \Rightarrow N(0,1) \tag{16}$$

with

$$t_h = \frac{\hat{\beta}_h}{\sqrt{(\frac{1}{T-h}\sum_{t=1}^{T-h}\hat{u}_t(h)^2)(\sum_{t=1}^{T-h}\underline{\varphi}_t^2)^{-1}}} \tag{17}$$

where $\hat{u}_{t+h}(h) = \Delta^h e_{t+h} - \hat{\alpha}_h - \hat{\beta}_h\varphi_t$ are the estimated residuals and $\underline{\varphi_t}$ is the demeaned value of φ_t.

Inference on $\hat{\beta}_h$ can thus be done by standardizing the standard t-statistic by the square root of the regression horizon h.

B Data Appendix

B.1 Commodities

Table A1. Commodity Description

Norgate Ticker	Bloomberg	Name	Exchange	Contract Months
Agricultural				
CC2	CC	Cocoa	NYSE Liffe	H,K,N,U,Z
KC2	KC	Coffee	ICE/NYBOT	H,K,N,U,Z
C-	C-	Corn	CBOT	H,K,N,U,Z
CT2	CT	Cotton	ICE/NYBOT	H,K,N,V,Z
FC	FC	Feeder cattle	CME	F,H,J,K,Q,U,V,X
LH	LH	Hogs	CME	G,J,K,M,N,Q,V,Z
LC	LC	Live cattle	CME	G,J,M,Q,V,Z
LB	LB	Lumber	CME	F,H,K,N,U,X
O-	O-	Oats	CBOT	H,K,N,U,Z
OJ2	JO	Orange Juice	ICE/NYBOT	F,H,K,N,U,Z
S2	S-	Soybeans	CBOT	F,H,K,N,U,X
BO2	BO	Soybean Oil	CBOT	F,H,K,N,Q,U,V,Z
SB2	SB	Sugar	ICE/NYBOT	H,K,N,V
W-	W-	Wheat	CBOT	H,K,N,U,Z
Metals				
GC	GC	Gold	NYMEX	G,J,M,Q,V,Z
PA2	PA	Palladium	NYMEX	H,M,U,Z
PL2	PL	Platinum	NYMEX	F,J,N,V
SI2	SI	Silver	NYMEX	H,K,N,U,Z
Energy				
CL	CL	Crude oil	NYMEX	F-Z
HO	HO	Heating oil	NYMEX	F-Z

Notes: Futures Contracts terminology: January = F, February = G, March = H, April = J, May = K, June = M, July = N, August = Q, September = U, October = V, November = X, December = Z

Table A2. Composition of Aggregate Convenience Yield Indices

	φ^{pc1}	φ^{pc2}	φ^{AUS}	φ^{CAN}	φ^{NZ}
Agricultural					
Cocoa	0.040	-0.261			
Coffee	0.141	-0.062			
Corn	0.183	-0.145		0.010	
Cotton	0.123	0.007	0.098		
Feeder Cattle	0.120	0.053	0.138	0.076	0.500
Hogs	0.048	0.001		0.035	
Live Cattle	0.128	0.053	0.138	0.076	0.500
Lumber	0.122	0.113		0.266	
Oats	0.063	-0.101			
Orange Juice	0.110	-0.047			
Soybeans	0.069	-0.125			
Soybean Oil	0.088	-0.240			
Sugar	-0.016	0.218	0.009		
Wheat	0.162	-0.016	0.290	0.067	
Metals					
Gold	0.112	0.283	0.328	0.045	
Palladium	0.157	0.019			
Platinum	0.112	0.210			
Silver	-0.065	0.230		0.006	
Energy					
Crude Oil	0.181	0.076		0.419	
Heating Oil	0.151	0.055			

Notes: This Table reports the weights used to compute aggregate convenience yield indices. The first two columns report the scoring coefficients used to calculate the principal component scores φ^{pc1} and φ^{pc2}. Columns 3 to 5 report the export-weighted country-specific averages of the cross-section of individual commodity convenience. The weights represent the share of a commodity in a country's total commodity exports scaled to have a sum equal to 1.

B.2 Macroeconomic fundamentals

The four macroeconomic fundamentals I consider are:

Money supply: $\Delta(m_t - m_t^*)$, where $m_t = \ln M_t$ and M_t is M1, OECD Main Economic Indicators (MEI), seasonally adjusted. For the United Kingdom, I predict M1 from M0, IFS line 19MC.ZF before October 1986.

Industrial production: $\Delta(y_t - y_t^*)$, where $y_t = \ln Y_t$ and Y_t is the industrial production index, taken from IFS, line 66CZF, except for Australia and New Zealand for which no monthly series is available. For these countries, I compute monthly observations from quarterly data (IFS, line 66) using the same procedure as in Molodtsova and Papell (2009).

Inflation rate: $\Delta(p_t - p_t^*)$, where $p_t = \ln P_t$ and P_t is the CPI price level from IFS, line 64 ZF. For the United Kingdom we take the price level series from OECD MEI until December 1987.

Interest rate: $i_t - i_t^*$, where i_t is the monthly return calculated from the money market rate, IFS line 60B.

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