Testing the Present-Value Model of the Exchange Rate with Commodity Currencies

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Abstract

Countries that specialize in commodity exports often exhibit a correlation between the relevant commodity price and the value of their currency. We explore an explanation for this correlation based on the present-value, monetary model of the exchange rate. An increase in the commodity price leads to an increase in the expected, future policy interest rate and so to an immediate appreciation. We test the model's over-identifying restrictions for Canada, Australia, and New Zealand. There, controlling for the effect of commodity prices in predicting current and future monetary policy leaves those prices no significant, remaining role in statistically explaining exchange rates.

JEL classification: F31, F41, E52.

Keywords: commodity currency, exchange rate, monetary policy

1. Introduction

Canada, Australia, and New Zealand often are described as having commodity currencies. For these countries there is an obvious correlation between an export commodity price (or an index of them) and the exchange rate. Commentators typically remark on this correlation at high frequency but it also is evident in the monthly or quarterly data that macroeconomists usually study. In fact, one might almost say there is an 'exchange-rate connect puzzle' for these countries: a reliable correlation between the nominal exchange rate and an exogenous variable, without a widely-accepted theory to explain that.

Meanwhile, macroeconomic models predict that an improvement in a country's terms of trade (for example through an increase in the prices of its export commodities) will lead to a *real* appreciation. An example is the MXN model described by Uribe and Schmitt-Grohé (2017, chapter 8). But theory also predicts that the form that this real appreciation takes will depend on monetary policy. For example, with a fixed nominal exchange rate there will be an increase in the price of nontraded goods and hence an increase in domestic inflation. In contrast, for a country that successfully targets inflation, one would expect the real appreciation to occur through a nominal appreciation. The three countries listed above also have inflation targeting in common, which thus suggests a simple resolution of the puzzle.

The mechanism we study is this: An increase in the export commodity price leads participants in the foreign exchange market to expect a tightening of domestic monetary policy relative to policy in the US (perhaps to stabilize the inflation rate). The exchange rate reacts immediately to the change in expected future policy, inducing a correlation with the commodity price. Thus the effect of the commodity price on the nominal exchange rate is intermediated by the reaction of monetary policy.

This combination of correlation and theory sets the stage for a test of the effect of relative monetary policy on the exchange rate. To allow for effects of monetary policy that occur either in the current period or are expected to occur in the future, the setting is the traditional, present-value model of the exchange rate. We measure relative monetary policy using the difference between the central bank's policy interest rate and the US

federal funds rate. Forecasts of future monetary policy of course cannot be observed directly, so the paper uses their projections on commodity prices. We then present a test of the hypothesis that those prices contribute to exchange-rate movements only through this mechanism.

For all three countries, controlling for the effect of commodity prices in predicting current and future monetary policy leaves them no significant, remaining role in explaining exchange rates. Based on this empirical evidence, we are not arguing that the entire adjustment in the real exchange rate occurs through the nominal exchange rate. We are arguing that the nominal appreciations (in response to commodity price changes) that we study can be explained through the response of current and expected future monetary policy.

2. Measurement

For commodity-exporting countries, researchers have documented a correlation between the nominal exchange rate and commodity prices measured with an index or else using the price of an individual export commodity such as oil. Key studies include those of Ferraro, Rogoff, and Rossi (2015) and Kohlscheen, Avalos, and Schrimpf (2017). This correlation also is documented for the real exchange rate by Chen and Rogoff (2003), Cashin, Céspedes, and Sahay (2004), and Chen and Lee (2018).

Our study focuses on Canada, Australia, and New Zealand. These are the same countries studied by Chen and Rogoff (2003). We focus on these three countries because (a) they have targeted inflation for most of the period we study, (b) they are small, open economies so that commodity prices are exogenous, and (c) they have long-standing and widely followed commodity export price indexes.

We label the commodity price X_t and its logarithm x_t. To measure this we use the national, commodity price indexes of central banks, where possible. Central banks say that they track these series so it seems natural to assume that participants in foreign exchange markets do so too. We thus use the commodity price indexes of the Bank of Canada and the Reserve Bank of Australia. The Reserve Bank of New Zealand does not publish such

an index, but a private bank (ANZ) has done so since 1986 and its index is widely tracked. Each series is at monthly frequency. We use the versions in USD then deflate by the US CPI. Our statistical tests involve both the overall index x_t and its main component, labelled $x_{m,t}$. This is the energy component for Canada, the base metals component for Australia, and the dairy products component for New Zealand.

The exchange rate, S_t (with log s_t), is the value of the local currency in USD, also monthly. Thus an increase is a domestic appreciation. We measure the stance of monetary policy relative to that in the US by the difference between the policy interest rate in the home country and the US interest rate: $d_t \equiv i_t - i_t^*$. The online appendix provides definitions and graphs of each series.

3. Present-Value Model

Our goal is to derive a test of the hypothesis that commodity prices may affect the nominal exchange rate through their effect on monetary policy. To do this, we adopt a present-value model in which the nominal exchange rate depends on current and expected future monetary policy, relative to policy in the US.

Our hypothesis has two components. First, commodity price movements sometimes lead to a reaction from monetary policy. Although there may be some persistence in the policy rate due to interest-rate smoothing, future monetary policy can be partly forecasted with commodity prices. Commodity prices are highly persistent, which may enhance their role in forecasts. Second, the nominal exchange rate responds to both current and expected future policy interest rates, so it reacts immediately to the commodity price index.

Recall that s_t is the log exchange rate, x_t the log, real commodity price, and $d_t \equiv i_t - i_t^*$ the differential in policy interest rates relative to the US. The exchange rate is determined by the simplest, present-value, monetary model of the exchange rate, with relative monetary policy measured by d_t:

$$s_t = \alpha(1 - \beta)d_t + \beta \mathrm{E}_t s_{t+1} = \alpha(1 - \beta)\mathrm{E}_t \sum_{j=0}^{\infty} \beta^j d_{t+j}, \tag{1}$$

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(with the transversality condition implicit). Here α is a scale factor and β is a discount factor. If d_t has a unit root then s_t inherits that, with cointegrating parameter α.

One environment that gives rise to this condition is the traditional, monetary model of the exchange rate in its flexible price version (and for simplicity omitting income terms or a risk premium in the UIP condition) as described by Engel and West (2005, section IIIA) or Engel, Mark and West (2007, section 1). Obstfeld and Rogoff (2003, equation 30) derive a similar condition from a model with real balances in the utility function. In these models $\beta = \lambda/(1+\lambda)$ where λ is the interest semi-elasticity of money demand. These models use the relative money supply as the fundamental and have a long-run classical property that an increase in the money supply depreciates the currency equi-proportionately. Our main amendment is to measure the stance of monetary policy using the policy interest rate, rather than the money supply, a change that also is characteristic of exchange-rate models that use Taylor rules. Thus α can be interpreted as the money supply semi-elasticity with respect to the policy interest rate: The value $-\alpha$ measures the extent to which an increase in d reduces the log relative money supply. We thus expect α to be positive and below find $\hat{\alpha}$ to be positive: An increase in expected future i, relative to US i^*, leads to a rise in s, an appreciation of the domestic currency.

Engel and West (2005) and Engel, Mark, and West (2007) also provide a range of empirical evidence on the present-value approach. They stress the need to allow for the endogeneity of monetary policy. We test for the possibility that monetary policy is expected to respond to the commodity price and so the current exchange rate does so too. An implication of present-value models is that the exchange rate should Granger-cause fundamentals. This present-value model thus also is consistent with Chen, Rogoff, and Rossi's (2010) finding that nominal exchange rates (for commodity currencies) help forecast commodity prices. And it also can be consistent with the exchange rate's following a random walk, as shown by Engel and West (2004).

One might suppose that policy reacts to x_t because it is an indicator of future inflation, which is the true target of monetary policy. Bernanke, Gertler, and Watson (1997) argued that effects of oil-price shocks on macroeconomic variables are intermediated through re-

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actions of monetary policy. Several studies have noted that a shock to inflation may lead to a nominal appreciation (in the short run only) because of a policy response that raises short-term interest rates. Clarida, Galí, and Gertler (2002) show that optimal policy has this feature in a new Keynesian, two-country model. Engel and West (2006) derive this mechanism in a model with Taylor rules. Clarida and Waldman (2008) find evidence that inflation surprises are associated with immediate exchange-rate appreciations for countries that target inflation.

We conjecture that the effect of x on d here is due to its effect on each central bank's forecasts of inflation and the output gap but we do not solve a complete model to examine that idea, in part because Australia and New Zealand do not measure the CPI at monthly frequency. Our test is also different from those in the studies above, because we can rely on the exogeneity of x. And it applies whatever the mechanism by which x_t affects d_t.

4. Estimates and Tests

Pre-tests reported by Devereux and Smith (2018) suggest that the variables contain unit roots, yet a survey of sources on present-value methods with nonstationary variables shows that there is no consensus on what method to use for estimation. We adopt a simple, two-step procedure. The first step follows Campbell and Shiller's (1987) application of the Granger-Engle two-step method. Define a new variable $y_t \equiv s_t - \alpha d_t$ and then rewrite the present value (1) as:

$$y_t = \alpha \beta E_t \Delta d_{t+1} + \beta E_t y_{t+1} = \alpha \mathrm{E}_t \sum_{j=1}^{\infty} \beta^j \Delta d_{t+j}. \tag{2}$$

If d_t is $I(1)$ then d_t and s_t are cointegrated and $s_t - \alpha d_t$ is $I(0)$ so the variables in this second present value are stationary, which facilitates inference.

We first estimate the cointegrating relationship between s_t and d_t (with coefficient $\tilde{\alpha}$) and generate $\tilde{y}_t = s_t - \tilde{\alpha} d_t$. (A constant term is included but not reported.) In this environment it is well known that coefficients in the levels regressions will be estimated super-consistently. It also is well known that there may be bias in such estimates if dynamics are omitted. We estimate $\tilde{\alpha}$ by fully modified ordinary least squares as developed

by Phillips and Hansen (1990). This involves corrections for endogeneity and serial correlation that reduce the bias in OLS estimation of the static regression. The online appendix derives the cointegrating relationships and shows how they identify α.

The second step then involves estimation with stationary variables by GMM. The transformed model implies that:

$$E\big[(\tilde{y}_t - \tilde{\alpha}\beta\Delta d_{t+1} - \beta\tilde{y}_{t+1})|\Delta x_t, ..., \Delta x_{t-n}\big] = 0. \tag{3}$$

We estimate the discount factor β by continuously updated GMM. Current and lagged values of Δx_t can be valid instruments either because that series itself has higher-order, autonomous dynamics, or because monetary policy reacts to those lags. With one parameter to estimate, adopting more than one instrument allows the familiar J-test (and adds precision to the estimate $\hat{\beta}$). If the over-identifying restrictions hold, then the difference-equation residuals are not correlated with the current and lagged commodity-price growth rates. We know that s_t is correlated with x_t (as that is the reason for this paper to exist), so this test assesses whether any correlation remains once we control for the fact that commodity prices predict current and expected future values of d_t.

Table 1 summarizes the results of our test of conditions (3). The first column lists the countries, the second column gives $\tilde{\alpha}$, and the third column gives n, the number of lags of Δx_t in the instrument set. We then report $\hat{\beta}$. Once we add lags the last column gives the J-test statistic and its p-value.

Estimates $\tilde{\alpha}$ are positive, with t-statistics of 2.84, 3.75, and 3.56 for Canada, Australia, and New Zealand respectively. Discount factor point estimates $\hat{\beta}$ are plausible. Then the main finding is easy to report: None of the sets of restrictions is rejected at conventional levels of significance. The smallest p-value is 0.24. Thus we cannot reject the hypothesis that commodity prices are correlated with the exchange rate because those prices forecast differential monetary policy.

However, it is possible that the central bank in each country reacts not to the overall index of export commodity prices but instead to its largest component *and* that the foreign exchange market expects it to do so. In that case, our test power in table 1 may be low

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because the overall index is a noisy measure of this component. To examine that possibility, we repeat the calculations in table 1 but now with $x_{m,t}$, measured by the energy component for Canada, the base metals component for Australia, and the dairy products component for New Zealand. The results are in table 2. Now the t-statistic for $\tilde{\alpha}$ for Canada is 1.4 and the lowest p-value for the J-test is 0.10 for Australia when $n = 4$. But overall the results are similar to those in table 1, as $\tilde{\alpha}$ is positive, $\hat{\beta}$ takes plausible values, and the J-test does not reject at conventional significance levels.

The results in table 2 are of course not independent of those in table 1, for the components in table 2 are central to the indexes in table 1. But even performing our test across separate commodity price indicators may make it difficult to control test size (equivalently raise the risk of false rejections), given the relatively small number (415–427) of monthly observations possible in this study. We hope that the simplicity of the method will allow researchers to apply it for other countries or time periods. Of course, we are not arguing that the findings will apply to all commodity-exporting countries: They depend on the local monetary policy.

5. Conclusion

We examine a natural but under-studied explanation for a correlation between a country's commodity price index x_t and its nominal exchange rate s_t: s_t is determined by current and expected future values of an indicator of relative monetary policy and that indicator reacts to x_t.

We study three countries selected according to the criteria that (a) they have long histories of floating exchange rates, mainly under inflation targeting and (b) they have widely-followed export commodity price indexes. Our hypothesis is that the central bank follows this index in setting monetary policy and so practitioners in the foreign exchange market react to it too. For Canada, Australia, and New Zealand, J-tests show that there is little correlation between the exchange rate and the commodity price once we control for the role of the commodity price in predicting measures of current and expected future monetary policy.

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Table 1: Estimates and Tests
(Export Commodity Price Indexes)

$$\tilde{y}_t \equiv s_t - \tilde{\alpha}d_t$$

$$E\left[\left(\tilde{y}_t - \tilde{\alpha}\beta\Delta d_{t+1} - \beta\tilde{y}_{t+1}\right)\big|\Delta x_t, ..., \Delta x_{t-n}\right] = 0$$

Country	$\tilde{\alpha}$ (se)	n	$\hat{\beta}$ (se)	$J\,(df)$ p
Canada	0.245 (0.086)	0	0.908 (0.047)	
		4	0.954 (0.021)	4.9 (4) 0.29
		12	0.975 (0.016)	11.9 (12) 0.45
Australia	0.135 (0.036)	0	0.967 (0.080)	
		4	1.02 (0.073)	2.3 (4) 0.68
		12	0.925 (0.069)	5.4 (12) 0.94
New Zealand	0.089 (0.025)	0	0.907 (0.065)	
		4	0.977 (0.031)	5.5 (4) 0.24
		12	1.01 (0.024)	8.3 (12) 0.76

Notes: x is the log, real commodity price index, s the log nominal exchange rate in USD, and d the policy interest-rate differential. T=415 from 1986:1 to 2020:7 for New Zealand and T=427 from 1985:1 to 2020:7 for Canada and Australia. The cointegrating coefficient $\tilde{\alpha}$ is estimated by FM-OLS with 3 lags. Estimation of β is by continuously updated GMM with n lags of Δx_t as instruments. Constants and an instrument of ones are included in each equation but not shown.

Table 2: Estimates and Tests
(Price Index Components)

$$\tilde{y}_t \equiv s_t - \tilde{\alpha}d_t$$

$$E\left[(\tilde{y}_t - \tilde{\alpha}\beta\Delta d_{t+1} - \beta\tilde{y}_{t+1})\big|\Delta x_{m,t}, ..., \Delta x_{m,t-n}\right] = 0$$

Country	$\tilde{\alpha}$ (se)	n	$\hat{\beta}$ (se)	$J\,(df)$ p
Canada	0.456 (0.329)	0	0.957 (0.028)	
		4	0.978 (0.013)	3.4 (4) 0.50
		12	0.989 (0.008)	9.1 (12) 0.69
Australia	0.200 (0.093)	0	0.832 (0.153)	
		4	0.959 (0.035)	7.9 (4) 0.10
		12	0.955 (0.040)	11.2 (12) 0.51
New Zealand	0.149 (0.057)	0	0.883 (0.088)	
		4	0.968 (0.029)	7.3 (4) 0.12
		12	1.00 (0.020)	9.2 (12) 0.68

Notes: x_m is a log, real commodity price index component: energy for Canada, base metals for Australia, and dairy products for New Zealand, s the log nominal exchange rate in USD, and d the policy interest-rate differential. $T=415$ from 1986:1 to 2020:7 for New Zealand and $T=427$ from 1985:1 to 2020:7 for Canada and Australia. The cointegrating coefficient $\tilde{\alpha}$ is estimated by FM-OLS with 3 lags. Estimation of β is by continuously updated GMM with n lags of $\Delta x_{m,t}$ as instruments. Constants and an instrument of ones are included in each equation but not shown.