International Reserve Holdings: Interest Rates Matter!

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Abstract

We argue that the literature on optimal international reserve holdings in an era of high capital mobility fails to find interest rates are strongly significant factor because of the endogeneity of interest rates and reserves under fixed exchange rate regimes. Using two stage least squares we control for this and regain statistical significance for interest rates.

Keywords: International reserves; Buffer stock model; Exchange rate regimes; Endogeneity.

JEL classifications: E42, F31, F41.

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1. Introduction

For reserve authorities interest rates represent the opportunity cost of holding international reserves. Flood and Marion (2002) provide an important contribution to the literature by estimating optimal reserve holdings for countries under various monetary regimes. They find that, in a world with high capital mobility, interest rates are weakly significant and not robust in determining reserve holdings. We argue that this is due to the endogeneity of interest rates and reserves under a fixed regime. We control for this endogeneity using two-stage least squares and find that interest rates are significant determinants in our specification.

2. The Model and Method

With some success, buffer stock (or, inventory) models have been employed to explain post World War II international reserve holdings[1]. Frenkel and Jovanovic (1981), henceforth FJ, successfully applied Miller and Orr's (1966) inventory-theoretic framework to explain international reserve holdings. This work was updated significantly by Flood and Marion (2002).

The buffer stock model argues that reserve authorities choose a level of reserves that minimizes two expected costs. The first is the opportunity cost of holding reserves since international reserves offer a lower return than other assets. The second is the adjustment cost of restocking that is incurred whenever reserves reach some lower bound. This is generally modeled by constant costs times the standard deviation of a Wiener process, capturing that higher volatility increases the probability of incurring the restocking cost given some lower bound of desired reserve holdings. This is usually represented as:

$$R_0 = \sqrt{\frac{C\sigma}{r^{1/2}}} \quad (1)$$

where R_0 is the optimal starting level of reserves, C is a country specific nominal constant (fixed adjustment cost), σ is the standard deviation of the Wiener increment in the reserves time-series process (volatility measure), and r is the opportunity cost of holding reserves (usually approximated by a country's government bond yield).

A simple log transformation of expression (1) yields the original econometric specification (and results) from FJ.

$$\begin{aligned} \ln R &= \beta_0 + \beta_1 \ln \sigma + \beta_2 \ln r + u \\ &= b_0^i + \underset{(0.110)}{0.505 \ln \sigma} - \underset{(0.149)}{0.279 \ln r} \quad . \end{aligned} \quad (2)$$

These estimates (OLS standard errors in parenthesis) result from a sample of twenty two developed countries over the period of 1971 – 1975. The theoretically predicted values from equation (1) are $\beta_1 = 0.50$ and $\beta_2 = -0.25$.

Flood and Marion (2002) argue that FJ's estimates, especially the volatility elasticity of reserves, are likely to be biased upward. Their basic criticism arises from the fact that the original FJ approach does not separate typical incremental volatility, i.e. the Wiener increment, from the relatively large upward restocking adjustments that occasionally take place. These periodic adjustments cause the distribution for the time series to be skewed and therefore potentially lead to upwardly biased coefficient estimates. To address this issue, Flood and Marion build on the theoretical work of Flood and Garber (1984), Flood and Marion (1999), and Flood and Jeanne (2000) to develop a proxy variable for reserve volatility, σ^{FM}, referred to as the shadow fundamentals volatility. The measure is based on the economic fundamentals that drive the shadow exchange rate to its upper bound at the same time that reserves are driven to their lower bound. For later reference, we use Flood and Marion's σ^{FM} in our regression analysis as well. Details on calculating this variable are included in the Data Appendix.

Flood and Marion contribute to FJ's analysis in two additional ways. The first is to include alternative scaling variables for robustness purposes. The second is to use a standardized measure for the opportunity cost of holding reserves. FJ used a government bond yield, r. Flood and Marion use $i=(1+r)/(1+r^*)$, where r and r^* denote the domestic and US bond yields, respectively. The resulting fixed effect econometric specification is

$$\ln(R/X) = \beta_0 + \beta_1 \ln(\sigma^{FM} / X) + \beta_2 \ln i + u \tag{3}$$

where X is a scaling variable representing one of the following: unity, GNP, the price level, nominal value of imports, or M2. Using data on 36 countries during a period of high capital mobility (1988 – 1997), Flood and Marion show that the result of reserves increasing with volatility is robust but the prediction that reserves decrease with increases in the opportunity cost is not. When other explanatory variables are considered, specifically controlling for the degree of exchange rate flexibility and economic openness[2], the already weak support for the significance of the opportunity cost variable disappears completely[3]. The problem with the approach of Flood and Marion is that they fail to account for the endogeneity of interest rates and international reserves under a fixed exchange rate regime[4]. Accordingly their estimates for the coefficient of the opportunity cost are biased downward and inconsistent.

By uncovered interest parity, domestic (or foreign) interest rate movements under a flexible exchange rate regime influence the nominal exchange rate's level and/or rate of change, but do not influence international reserve holdings. That is, outside of optimal portfolio holding considerations, the two are not related.

Under a credibly fixed regime, however, any changes in the currency's level or expected rate of depreciation should only come about as a conscious policy decision. This requires foreign exchange sales or purchases of the domestic currency and thus a change in foreign reserves. Since this alters the quantity of domestic currency in private hands, it also alters domestic interest rates. Therefore, the change in interest rates is directly linked to changes in reserves. The two are endogenously determined in fixed regimes[5].

To address this endogeneity issue, we introduce a two-stage least squares estimation (2SLS) where the opportunity cost variable, $i = (1+r)/(1+r^*)$, is regressed against exogenous variables and then used in equation (3) as an estimate for i. The results show that after controlling for endoegneity issues, the opportunity cost of holding reserves is non-negligible and therefore it plays an important role in determining the level of reserve holdings, which is consistent with the findings of Grimmes (1993).

3. Data and Results

All data come from the IMF's International Financial Statistics. The specific codes series are in the Data Appendix by country. We first replicate some of the relevant results in Flood and Marion[6]. We then run 2SLS. All results are reported in Table 1.

Panel 1 of the table reports results obtained when only σ^{FM} and i are used as control variables. The estimated coefficients are close in magnitude and equal in sign to those reported by Flood and Marion (bottom section of Panel 1). There is strong evidence for a positive relationship between the level of reserve holdings and σ^{FM}. There is also some evidence for the negative relationship with i[7]. When additional control variables are introduced, Panel 2, the statistical significance of the opportunity cost disappears completely. This is the result we find contentious on endogeneity grounds.

[INSERT TABLE 1]

The instruments used in the first stage of our 2SLS are i_{t-1} and σ_{t-1}^{FM} (results in the Regression Appendix). We include a dummy variable, D, equal to one for a fixed exchange rate regime and zero otherwise. The regime dummy variable and the interaction term in the first stage regression clean out the effect of changes in the interest rate resulting from changes in reserves and the possible differences (intercept and slope) that arise due to a fixed exchange rate regime[8]. The adjusted R^2 is around 0.78 for all four of the weights considered, suggesting a reasonable fit for the estimation of the opportunity cost.

Panel 3 of Table 1 presents the second stage regression results. The sign and statistical significance of the coefficient for the variability of reserves are preserved. As in Flood and Marion, the magnitude of this coefficient is lower than in Panel 1 but similar to Panel 2. The opportunity cost coefficient, however, is always negative and statistically significant, supporting the theoretical prediction of the "buffer stock" model. Furthermore, the magnitude of the coefficient is not statistically different from the predicted value of -0.25[9]. The coefficients for capital flows and trade flows are always significant and positive, as predicted by economic theory.

4. Conclusion

We contribute to a line of literature that estimates the optimal reserve holdings for countries under various monetary regimes. Due to an endogeneity problem between interest rates and international reserve holdings under fixed exchange rate regimes, past work lost much of the significance and robustness of interest rates in determining reserve holdings. We control for this endogeneity using two-stage least squares and regain strong interest rate significance.

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Table 1.

Weighted by	No Weights	GDP Deflator	GNP	Imports
Panel 1				
Current Study Results				
σ(FM)	0.1743 *	0.1366 *	0.0807 *	0.0635 *
	0.0108	*0.0080*	*0.0079*	*0.0075*
i = (1+r)/(1+r*)	-0.1381 **	-0.1430 *	-0.0993 ***	0.0330
	0.0599	*0.0509*	*0.0512*	*0.0336*
Adj . R 2	0.88	0.900	0.780	0.770
Observations	472	472	467	472
Cross Sections	36	36	36	36
As Reported by Flood and Marion (2002)				
σ(FM)	0.1720 *	0.1417 *	0.1058 *	0.0908 *
	0.0344	*0.0264*	*0.0185*	*0.0203*
i = (1+r)/(1+r*)	-0.1421 ***	-0.1574 **	-0.0634	0.0134
	0.0855	*0.0744*	*0.0474*	*0.0496*
Adj. R2	0.904	0.919	0.839	0.823
Panel 2				
σ(FM)	0.0626 *	0.0628 *	0.0624 *	0.0602 *
	0.0065	*0.0065*	*0.0067*	*0.0061*
i = (1+r)/(1+r*)	-0.0432	-0.0459	-0.0501	-0.0300
	0.0413	*0.0406*	*0.0423*	*0.0423*
Total Trade/1	0.8639 *	0.8498 *	0.8061 *	0.6050 *
	0.0545	*0.0584*	*0.1095*	*0.1795*
Capital Flows/1	0.1070 *	0.1107 *	0.1215 *	0.1164 *
	0.0317	*0.0320*	*0.0343*	*0.0338*
Adj . R 2	0.940	0.938	0.822	0.779
Observations	472	472	467	472
Cross Sections	36	36	36	36
Panel 3				
σ(FM)	0.0617 *	0.0612 *	0.0592 *	0.0596 *
	0.0154	*0.0154*	*0.0123*	*0.0157*
i = (1+r)/(1+r*)	-0.1274 ***	-0.1430 **	-0.1571 **	-0.1303 ***
	0.0766	*0.0771*	*0.0822*	*0.0798*
Total Trade	0.8397 *	0.8134 *	0.6968 *	0.7095 *
	0.0694	*0.0806*	*0.1240*	*0.2656*
Capital Flows	0.1142 *	0.1177 *	0.1278 *	0.1186 *
	0.0348	*0.0345*	*0.0393*	*0.0340*
Adj . R 2	0.942	0.940	0.825	0.769
Observations	420	420	415	420
Cross Sections	36	36	36	36

* 1%, ** 5%, and *** 10%

Heteroskedasticity corrected standard errors for the estimated coefficients are reported in italics.

1/ Flood and Marion (2002) report that the estimated coefficients for these variables are positive and significant, but they do not elaborate on their magnitude.

Hansen's J-Statistic was used to verify the validity of the overidentifying restrictions for the 2SLS

Data Appendix

All data are from the IMF's International Financial Statistics Database [codes].

International Reserves: [1.SZF] in dollars using $/SDR exchange rate [111.AA.ZF].

US GDP deflator: [11199BIRZF] used in computing real international reserves.

Total Trade: Exports [70.DZF] plus Imports [71.DZF].

Capital Flows: Sum of capital inflows, [78BEDZF, 78BGDZF, 78BIDZF], and capital outflows, [78BDDZF, 78BFDZF, 78BHDZF].

Shadow Fundamentals volatility, σ^{FM}: Following Flood and Marion, the shadow fundamental rate is $D/(P^*\gamma)$, where D, P^*, and γ denote Domestic Credit [32..ZF] and the Foreign Price level (i.e. US price level) [11164..ZF], and a constant[10]. The volatility measure is generated by computing the standard deviation over the previous fifteen years of the trend-adjusted annual changes in the shadow fundamental rate.

Countries, interest rates, and GNI/GDP series:

Country Name	Interest Rate	Gross National Income/Gross Domestic Product	Country Name	Interest Rate	Gross National Income/Gross Domestic Product
ARGENTINA	60L.ZF	99A.ZF	JAMAICA	60C.ZF	99A.ZF
AUSTRALIA	61.ZF	99A.CZF	JAPAN	60.ZF	99A.CZF
AUSTRIA	60.ZF	99A.ZF	KOREA	60.ZF	99A.ZF
BELGIUM	60C.ZF	99A.ZF	MALAYSIA	60L.ZF	99A.ZF
BRAZIL	60B.ZF	99A.ZF	MEXICO	60L.ZF	99B.CZF
CANADA	60P.ZF	99A.CZF	NETHERLANDS	61.ZF	99A.ZF
CHILE	60P.ZF	99A.ZF	NEW ZEALAND	61.ZF	99A.CZF
COLOMBIA	60.ZF	99A.ZF	NORWAY	61.ZF	99A.ZF
COSTA RICA	60.ZF	99A.ZF	PAKISTAN	60.ZF	99A.ZF
DENMARK	60.ZF	99A.ZF	PHILIPPINES	60.ZF	99A.ZF
FINLAND	60.ZF	99A.ZF	PORTUGAL	60.ZF	99B.ZF
FRANCE	60P.ZF	99A.CZF	SOUTH AFRICA	60P.ZF	99A.CZF
GERMANY	60B.ZF	99A.CZF	SPAIN	60.ZF	99A.ZF
INDIA	60B.ZF	99A.ZF	SWEDEN	60.ZF	99A.ZF
INDONESIA	60L.ZF	99A.ZF	SWITZERLAND	61.ZF	99A.ZF
IRELAND	60.ZF	99A.ZF	THAILAND	60P.ZF	99A.ZF
ISRAEL	60P.ZF	99A.ZF	UNITED KINGDOM	60C.ZF	99A.CZF
ITALY	61.ZF	99A.CZF	VENEZUELA, REP. BOL.	60.ZF	99A.ZF

Regression Appendix

Weighted by	No Weights	GDP Deflator	GNP	Imports
First Stage Regression				
constant	-0.5310	-0.7519 **	-0.9278 *	-1.0645 *
	0.4082	*0.3327*	*0.1527*	*0.1489*
i (t-1)	0.5590 *	0.5578 *	0.5604 *	0.5558 *
	0.0326	*0.0328*	*0.0331*	*0.0328*
D	0.0085	0.1192 **	0.1645	0.2151 **
	0.0384	*0.0500*	*0.1128*	*0.1004*
σ(FM)	0.0271	0.0236	0.0342 *	0.0351 ***
	0.0192	*0.0193*	*0.0189*	*0.0188*
σ(FM) (t-1)	-0.1160 *	-0.1210 *	-0.1265 *	-0.1268 *
	0.0166	*0.0169*	*0.0162*	*0.0161*
D σ(FM)	0.0279 **	0.0255 **	0.0124	0.0196 ***
	0.0119	*0.0120*	*0.0108*	*0.0109*
D σ(FM) (t-1)	-0.0185 ***	-0.0079	-0.0011	-0.0018
	0.0102	*0.0079*	*0.0032*	*0.0035*
Total Trade	0.0439	0.0518	0.0464	0.3102 ***
	0.0462	*0.0537*	*0.0824*	*0.1680*
Capital Flows	0.0162	0.0136	0.0104	0.0019
	0.0228	*0.0226*	*0.0231*	*0.0221*
Adj . R 2	0.785	0.7838	0.7806	0.786

* 1%, ** 5%, and *** 10%

Heteroskedasticity corrected standard errors for the estimated coefficients are reported in italics.
Durbin Watson statistic shows now evidence of serial correlation.
(t-1) indicates the variable is lagged one period.

[1] Bahmani-Oskooee and Brown (2002) place the buffer stock model in its proper context in the relevant literature while also providing an excellent review of the literature in general.

[2] The measure of exchange rate flexibility is computed by using the standard deviation of the innovation to the percentage change in the nominal effective exchange rate. Openness is captured by capital and trade flows.

[3] Sebastian Edwards (1985) found that interest rates were significant. Flood and Marion, however, emphasize that their results are for the modern era of high capital mobility. Since Edwards (1985) was prior to the high capital mobility era, Flood and Marion's results remain the ones we must address.

[4] Edwards (1983) and Bahmani-Oskooee (1987) show that the degree of exchange rate flexibility plays an important role in determining the demand for international reserves as well as the adjustment between the actual and the desired level.

[5] The relationship still holds if the fixed regime does not enjoy perfect credibility or occasionally sterilizes. Perfect non-credibility or perpetual sterilization implies that "fix" is a misclassification. Any thing less than that adds noise to the empirical observation of the relationship but does not eliminate the endogenous relationship itself.

[6] The sample period considered in FM is 1988 – 1997, while the one considered in this study is 1987 – 2000. We use the same 36 countries (listed in the Data Appendix) as Flood and Marion.

[7] As in Flood and Marion, the estimated coefficient is actually positive but statistically insignificant when the level of imports is used as the scale variable.

[8] The exchange rate regime classification follows Levy-Yeyati and Sturzenger (2002).

[9] For two of the four cases the coefficient is not statistically different from -0.25 at the five percent confidence level, while for the other two this holds at the ten percent level.

[10] Further details on the shadow fundamental rate are in Flood and Marion (1999).