ESSAYS IN EXCHANGE RATES AND INTERNATIONAL FINANCE

A thesis submitted for the Degree of Doctor of Philosophy

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ABSTRACT

This thesis pertains to international finance and models of exchange rate determination as well as efficiency of the market for foreign currency. The first chapter is an introduction where we discuss the advent of flexible exchange rate regimes and the development of monetary models of exchange rate determination as well as present a framework for this thesis. In the second chapter we consider the historical failure of monetary models of the exchange rate and revisit the standard real interest differential (RID) model (Frankel, 1979a). The Great British Pound (GBP) and Canadian Dollar (CAD) vis-à-vis the United States dollar (USD) are examined during the period 1980:Q1 - 2015:Q4, a time characterized by flexible exchange rate regimes and heightened capital mobility across borders. Unit root properties of the sample variables are examined and the Johansen (1995) methodology is applied to test for cointegration. The RID model yields a single cointegrating relation however tests of long-run exclusion (LE) and weak exogeneity (WE) show that the RID model is not a coherent model of the GBP and CAD against the USD. The study is furthered by examination of the hybrid monetary model (Hunter and Ali, 2014). The hybrid model is tested for comparison with Japan, as the post 2007-2009 financial crisis period is branded by zero-lower bound interest rates, a phenomenon first experienced by Japan for any prolonged period of time. The hybrid model in addition yields a single relation however tests of LE and WE show that the long-run projection is reversed and that a coherent relationship exists between the GBP and CAD vis-à-vis the USD and variables related to monetary fundamentals as well as long-run economic activity.

In the third chapter we examine efficiency of the market for foreign currency. The lead-lag pricing relationship between spot and futures rates is discussed and a panel employing data for the GBP, Australia Dollar (AUD), CAD, Brazilian Real (BRL) and South

African Rand (ZAR) vis-à-vis the USD is constructed at several intervals prior to expiry. The Johansen (1995) methodology is applied and shows that spot and futures rates cointegrate and that the cointegrating vector is the basis. Unit root properties for the basis are also examined and found to be integrated of order one or I(1). We therefore show that the market for foreign currency functions efficiently and that profitable arbitrage opportunities exist that restore prices to parity levels. This study is of particular significance in view of the markets' growing share and need for greater transparency to lay down appropriate regulation that limits systematic risk.

In the fourth chapter we re-examine monetary models of the exchange rate and consider the USD vis-à-vis the Japanese Yen (JPY) in view of the Japanese economy's slow growth in the post 2007-2009 financial crisis period. We test the standard RID monetary model as a framework for modelling the USD/JPY exchange rate however tests of WE show that the nominal exchange rate is weakly exogenous so drives the system instead of adapting to it. The hybrid monetary model developed by Hunter and Ali (2014) is adjusted in consideration of the current period of sluggish economic growth in Japan by incorporating differentials related to traded and non-traded goods productivity (Rogoff, 1992). The adjusted hybrid model produces a single cointegrating relation and joint tests of LE and WE show that the nominal exchange rate cannot be long-run excluded and is not weakly exogenous so that the adjusted hybrid model is a coherent long-run model of the USD/JPY nominal exchange rate.

In the fifth chapter we conclude and summarize the findings of the three studies presented in this thesis as well as provide practical recommendations for further study such as construction of dynamic error correction models and assessing out-of-sample forecasting performance for the extended monetary models examined in chapters two and four.

Further development of the study for effectively functioning foreign exchange markets as presented in chapter three is in addition discussed in the final chapter.

We contribute to the extant literature by showing in chapter two that the conventional RID monetary model of the exchange rate for the GBP and CAD vis-à-vis the USD can be rejected. A single econometric specification can be adapted to explain the long-run exchange rate for the GBP/USD exchange rate while an extended model is effective in providing an explanation of the long-run CAD/USD exchange rate. In chapter three we demonstrate that the spot and futures markets for five bilateral exchange rates function effectively across developed and developing countries. Lastly, we show in Chapter four that the model of the USD/JPY exchange rate due to Hunter and Ali (2014) appears a specific case and that the USD/JPY is not readily distinguished from a random walk in the context of a monetary model that considers traded and non-traded goods productivity differentials.

Dedicated to my mother without whom, I could never have done this

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DECLARATION

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Acronyms A to I	
A	
ADF	Augmented Dickey-Fuller test
ADF-GLS	Augmented Dickey-Fuller – Elliot, Rothenberg and Stock test
ARCH	autoregressive heteroskedasticity
ARDL	autoregressive distributed lag
AUD	Australian Dollar
B	
BIS	Bank for International Settlements
BRL	Brazilian Real
C	
CAD	Canadian Dollar
CME	Chicago Mercantile Exchange
CPI	consumer price index
CSD	Cross sectional dependence
D	
DF-GLS	Dickey-Fuller – Elliot, Rothenberg and Stock test
DM	Deutsche Mark
E	
ESTAR	Exponentially smooth transition autoregressive model
EGARCH	Exponential generalized autoregressive conditional heteroskedasticity
F	
FRF	French Franc
FPM	Flexible price model
G	
G7	group of the seven most advanced economies in the world including America, Canada, Japan, Great Britain, Germany, France and Italy
GARCH	Generalized autoregressive conditional heteroskedasticity
GBP	British Pound
GDP	gross domestic product
GLS	generalized least squares
GMM	generalized methods of moments
GNP	gross national product
H	
HBS	Harrod-Balassa-Samuelson effect
I	
I (0)	level stationary
I(1)	integrated of order one
ICE	International Currency Exchange
IDR	Indonesian rupiah
IFS	International Financial Statistics
IMF	International Monetary Fund
INR	Indian Rupee
IPI	industrial price index
ITL	Italian Lira

Acronyms J to Z	
J	
JPY	Japanese Yen
K	
KRW	South Korean won
L	
LE	long-run exclusion
M	
M1	money supply - the sum of currency in circulation and overnight deposits
M2	the sum of M1, deposits with an agreed maturity of up to two years and deposits redeemable at notice of up to three months
MEI	Main Economic Indicators
MFPM	flexible price monetary model
MRID	modified real interest differential model
MSM	Markov switching model
MSWARCH	Markov switching autoregressive conditional heteroscedasticity
O	
OECD	Organization of Economic Co-operation
OLS	ordinary least squares
P	
PPI	production price index
PPP	purchasing power parity
R	
RID	real interest differential
RM	Malaysian Ringitt
RMSE	root-mean-square error
S	
S	stationarity
SBIC	Schwarz Bayesian information criterion
SE	strict exogeneity
T	
THB	Thai baht
U	
UIP	uncovered interest parity
UK	United Kingdom
USA	United States of America
USD	United Stated Dollar
V	
VAR	vector autoregression
VECM	vector error correction model
W	
WE	weak exogeneity
WPI	wholesale price index
WTI	West Texas Intermediate
Z	
ZAR	South African Rand

CHAPTER ONE

INTRODUCTION

International investment and heightened capital mobility, since the collapse of the Bretton Woods regime in the early 1970s, have amplified the scope for financial yield while decreasing the extent of trading liability (Frankel, 1991; Cavallo, 2006; Chinn, 2006; Chinn and Ito, 2007; Batten and Szilagyi, 2010). A world that is economically consolidated presents the advantage of borrowing and lending across borders so that consumption is smoothed all through the business cycle (Obstfeld and Rogoff, 1995). The possibility of magnified returns on investment abroad is teamed with diverse liabilities including risk that is country specific and more precisely exchange rate risk. In the following section we discuss models of exchange rate determination as a backdrop to the subsequent three analyses in this thesis.

Establishment of the flexible exchange rate regime has necessitated greater depth of understanding of the exchange rate market. A plethora of literature exists concerning models of exchange rate determination. Leading models relate to the monetary base or assets and include the flexible-price (Bilson, 1978; Frenkel, 1976; Kouri, 1976 and Mussa, 1976 and 1979) sticky-price (Dornbusch, 1976; Frankel, 1979a; Buiter and Miller, 1981) and real interest differential (RID) models (Frankel, 1979a) while another branch of models pertain to the mobility of goods or portfolio balance (Kouri, 1976; Branson, 1977, 1983a and 1984; Frankel, 1979b, Allen and Kenen, 1980; Dornbusch and Fischer, 1980; Isard, 1983). Monetary models of the exchange rate build a relationship between two countries' anticipated inflation or price levels and the exchange rate. Despite the abundant literature a dearth of successful models of exchange rate determination exist.

Monetary models of exchange rate determination have been rejected based on biased and inconsistent coefficients (Driskill and Sheffrin, 1981), coefficient estimates that are insignificant or of a sign negating theory (Dornbusch, 1980; Haynes and Stone, 1981 and Frankel, 1984a) as well as the absence of cointegrating relationships (Baillie and Selover, 1987; Papell, 1997; Meese and Rogoff, 1988; Boothe and Glassman, 1987; McNown and Wallace, 1989; Baillie and Pecchenino, 1991), the presence of residual autocorrelation (Branson, Halttunen, and Masson, 1977; Lewis, 1988; Frankel, 1993), regime shifts or structural breaks (Goldberg and Frydman, 1996) and erratic expectations (Haache and Townend, 1981). It has in addition been shown that the out-of-sample naïve random walk model overrides long-run structural models of the exchange rate in most instances (Meese and Rogoff, 1983a; Chinn and Meese, 1995; Groen, 2005).

Monetary models have however yielded some success upon application of the Johansen (1995) methodology which is an extension of the Engle and Granger (1987) two-step procedure for identifying long-run relations (MacDonald and Taylor, 1994; McNown and Wallace, 1994; Moosa, 1994; Cushman et al., 1996; Choudhry and Lawler, 1997; Kouretas, 1997; Diamandis et al., 1998; Cushman, 2000; Tawadros, 2001 and Francis et al., 2001). Success has also been garnered upon application of the Johansen (1995) methodology in a multivariate framework (MacDonald and Taylor, 1991) and upon conditioning for regime shifts (Sarno, Valente and Wohar, 2004).

This thesis comprises three studies pertaining to exchange rates and international finance. In the first study, presented in chapter 2, we re-examine the traditional RID monetary model (Frankel, 1979a) and the hybrid monetary model (Hunter and Ali, 2014) for the GBP and CAD vis-à-vis the USD during the post Bretton Woods period 1980:Q1 – 2015:Q4. The time frame related to the United States, United Kingdom and Canada is presently linked to historically lower bound interest rates. In the second study, presented in chapter 3, we examine the efficiency of the market for foreign currency and the theory that the market for futures leads the cash market so that deviations from a theoretical futures value are suggestive of the presence of excess returns and arbitrage opportunities that re-establish equilibrium price levels. In the third study, presented in chapter 4, a monetary model of the USD/JPY exchange rate is developed that incorporates the instrumental variables employed by Rogoff (1992) in a fixed factor neoclassical model formulated to examine consumption smoothing.

In the first study, presented in chapter 2, we examine two monetary models of the GBP and CAD vis-à-vis the USD. The traditional RID monetary model (Frankel, 1979a) and hybrid monetary model (Hunter and Ali, 2014) are revisited to identify a meaningful long-run relationship between the exchange rate and variables related to the monetary base as well as real economic activity. The Johansen (1995) methodology is applied to quarterly data for the period 1980:Q1-2015:Q4 that is marked by flexible exchange rate regimes and a more recent period of prolonged zero lower bound exchange rates. Time series properties of the variables in both models are first examined as the series under investigation are required to be first difference stationary or I(1) as the Johansen (1995) methodology can only establish a long-run relationship amongst variables that are integrated of order one or I(1). Historical failure of the traditional RID model has been attributed to unstable money demand functions and the breakdown of PPP so to address these shortcomings variables related to the real economic architecture are incorporated in the conventional RID model to form the hybrid monetary model. Trace test statistics for both models identify a single cointegrating relation however several resulting variable coefficients for the RID models are of the wrong sign and magnitude. Furthermore, tests of LE and WE are conducted to ascertain model coherency. The findings resonate with Hunter and Ali (2014) as a single cointegrating relation is identified for the hybrid models and the exchange rate is not weakly exogenous and cannot be LE so is driven by the cointegrating framework. Success of the hybrid model highlights the significance of variables related to long-run economic activity and expectations.

In the second study, presented in chapter 3, we investigate the relationship between spot and futures rates in the market for foreign currency. Spot prices are likely to be influenced by the behaviour of market participants who act in accordance with information and expectations about the future in both the spot and futures market. A framework for modelling the pricing relationship is described by evaluation of the lead-lag pricing procedure as well as the procedure linked to deviations from hypothesized futures value. The pricing relationship between the spot and futures contracts for the GBP, CAD, AUD, BRL and ZAR vis-à-vis the USD is subsequently modelled and the hypothesis that the futures market leads the market for spot is tested. The pricing relationship is indicative of market efficiency and the way the market for foreign currency internalizes information. Deviations from theoretical futures values suggest the existence of excess returns and opportunities for arbitrage that return prices to equilibrium levels. To test the pricing relationship spot and futures prices are examined at different intervals; prior to maturity and at maturity. All spot rates and futures prices for the currency pairs in our sample are first difference stationary and tests of cointegration identify a single cointegrating vector for all five models. Stochastic properties of the basis are subsequently investigated, and it is shown that the basis for all five currency pairs is first difference stationary. The stochastic properties of the basis are further tested in a panel setting and confirm that the null of stationarity cannot be rejected so the basis is indeed stationary. The findings illustrate that the spot and futures market for the exchange rates in our sample is efficient therefore the futures market delivers risk free arbitrage opportunities and the spot market achieves its purpose as an appropriator of means.

In the third study, presented in chapter 4, we revisit monetary models of the exchange rate but focus on consumption smoothing by incorporating variables related to traded and non-traded goods productivity and terms of trade. Here the fixed-factor neoclassical model developed by Rogoff (1992) to examine the random walk behaviour of the real USD/JPY exchange rate is discussed as well as the Harrod-Balassa-Samuelson (HBS) effect. The relationship between the real USD/JPY exchange rate and individual variables related to consumption smoothing namely productivity in manufacturing, government consumption expenditure and the real oil price is examined by means of the Johansen (1995) test for cointegration. Tests of cointegration confirm a relationship between the real USD/JPY exchange rate and variables related to traded and non-traded goods productivity. We advance the study by estimating the standard RID monetary model for the USD/JPY nominal exchange rate. The RID model in view of the estimated coefficients fails to establish a meaningful long-run relationship between monetary fundamentals and the nominal exchange rate. In addition, results of tests for LE and WE show that the exchange rate operates like a random walk so cannot predict fundamentals that drive the framework represented in the model. Since the RID model is not an empirical model of the exchange rate the study is furthered by expanding the traditional RID model to incorporate factors related to traded and non-traded goods consumption, terms of trade and share prices. The adjusted hybrid model yields a single cointegrating relationship and joint tests of LE and WE confirm that the system drives the nominal USD/JPY exchange rate. The adjusted hybrid monetary model is therefore a coherent empirical model.

In Chapter 5 we present our conclusions and discuss the contribution of this thesis to the extent literature and provide some practical recommendations for further studies. We conclude that PPP is no longer a driver of the exchange rate and note the importance of financial assets that drive the system and advise that an error correction model might be constructed to further investigate the short-run dynamics among the variables considered in the RID model and hybrid monetary model investigated in chapters 2 and 4. The most important observation gained from the first and third studies is that prices do not drive the exchange rate in a world where financial assets exhibit substantial influence. We in addition recommend that the model for effective functioning of foreign exchange markets examined in the second study, presented in chapter 3, might be extended to stock markets and might include volatility by employing the multivariate generalized autoregressive conditional heteroskedastisity (GARCH) model considered by Bollerslev et al. (1988). We note that the model developed in the second study is particularly apt for examining market efficiency and can be adapted to other markets to further comprehend market functions and increase market transparency.

CHAPTER TWO

MONETARY MODELS OF THE GREAT GBP AND CAD VIS-À-VIS THE US DOLLAR: AN EMPIRICAL INVESTIGATION

2.1. Introduction

Abandonment of the fixed exchange rate regime in the early 1970's stimulated development of models of exchange rate determination and widened the scope of study of exchange rate behaviour. Foremost models pertain to monetary fundamentals or assets and comprise the flexible-price (Bilson, 1978; Frenkel, 1976; Kouri, 1976 and Mussa, 1976 and 1979) sticky-price (Dornbusch, 1976; Frankel, 1979a and Buiter and Miller, 1981) and real interest differential (RID) models (Frankel, 1979a) while a further stream of models focuses on the flow of goods or portfolio balance (Kouri, 1976; Branson, 1977, 1983a and 1984; Frankel, 1979b, Allen and Kenen, 1980; Dornbusch and Fischer, 1980 and Isard, 1983).

Monetary models assume complete markets where transaction costs are negligible and establish a link between two countries expected inflation or price levels and the exchange rate. Exchange rates move in tandem with the difference between prices on domestic and foreign goods that in complete markets are identical in every respect aside from currency of denomination. A higher price on a domestic good would result in an appreciation of the domestic currency and hence stimulate capital inflows. This link between exchange rate movement and price levels is integral to the establishment of monetary policy. Monetary variables are central to models of exchange rate determination however their inclusion relished no more than initial success (Frenkel, 1976; Bilson, 1978; Hodrick, 1978; Wilson, 1979; Hooper and Morton, 1982). Merger of data from the late 1970s and onwards led to rejection of the model on the grounds of biased and inconsistent coefficients (Driskill and Sheffrin, 1981), insignificant variable estimates or estimates of a sign opposing theory (Dornbusch, 1980; Haynes and Stone, 1981 and Frankel, 1984) and the absence of cointegrating relationships (Baillie and Selover, 1987; Papell, 1997; Meese and Rogoff, 1988; Boothe and Glassman, 1987; McNown and Wallace, 1989; Baillie and Pecchenino, 1991). The rejection of monetary models is explained by the absence of an empirical link between money supply and the exchange rate; and the presence of residual autocorrelation (Branson, Halttunen, and Masson, 1977; Lewis, 1988; Frankel, 1993).

To identify a long-run cointegrating relationship, Johansen (1995) and Phillips (1991) expanded upon Engle and Grangers' (1987) two-step procedure for identifying long-run relations. Subsequently the Johansen (1995) methodology has been employed to identify long-run cointegrating relationships in various studies related to models of exchange rate determination[1] (e.g. MacDonald and Taylor, 1994; McNown and Wallace, 1994; Moosa, 1994; Cushman et al., 1996; Choudhry and Lawler, 1997; Kouretas, 1997; Diamandis et al., 1998; Cushman, 2000; Tawadros, 2001 and Francis et al., 2001;) as well as in our study of the traditional RID monetary model and an augmented RID model presented in section four of this chapter. Notwithstanding, success of the procedure for identifying a long-run equilibrium, model variable estimates have, more frequently than not, been reported to be of a sign opposing theory and of the incorrect magnitude (Driskill and Sheffrin, 1981; Dornbusch, 1980; Haynes and Stone, 1981 and Frankel, 1984a). Widespread rejection of monetary models is attributed to the inadequacy of monetary fundamentals to describe fluctuations in the currency market.[2]

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[1] MacDonald and Taylor (1991) apply the Johansen (1995) methodology in a multivariate framework to the USD/DM, USD/GBP and USD/JPY and are unable to reject the existence of a long-run cointegrating relationship on the basis of identified underlying features of the economy that drive these rates.

[2] Goldberg and Frydman (1996) ascribe failure of monetary models to regime shifts or structural breaks and erratic expectations. Several studies have shown that upon conditioning that upon conditioning for regime shifts (Sarno, Valente and Wohar, 2004) the out-of-sample structural model beats the random walk by significant margins in root-mean-square error (RMSE) (MacDonald and Taylor, 1994 and Macdonald and Marsh, 1997).

Hunter (1992) re-examines the USD/GBP effective exchange rate and tests for cointegration between purchasing power parity (PPP) and uncovered interest parity (UIP) whereby he confirms the existence of such a relationship by incorporation of a global oil price as well as other prices and interest rates for both countries in the model. Hunter (1992) confirms by means of a multivariate cointegration technique that an unrestricted monetary model is a well-founded system for examining long-run exchange rate behaviour and that the unrestricted model beats the out-of-sample random walk model.[3]

PPP is unlikely to hold in the age of flexible exchange rate regimes where capital flows are soaring and money demand remains unstable. Since PPP and a stable money demand function are cornerstones of the monetary model it too is doomed to fail in the current era. In pursuit of constructing a predictable model for the equilibrium exchange rate it would be rational to incorporate real factors that can be visibly modelled to reflect a meaningful long-run relationship between the nominal exchange rate and monetary fundamentals. By incorporation of real dynamics inherent to monetary operations and currency markets one might hope to achieve coefficients consistent with theory and supersede overidentifying restrictions imposed by the assumption of rational expectations (Driskill and Sheffrin, 1981).

Friedman (1988) explains that stock prices affect money demand by a positive wealth effect where the demand for money grows in tandem with increased nominal wealth as well as heightened stock market transactions. Conversely the negative substitution effect

[3] Meese and Rogoff (1983a) examine the link between the USD/GBP exchange rate and various monetary bases from 1976 to 1981 by application of three structural models and a VAR to account for simultaneity bias however neither outperforms the random walk thus it is inferred that the models fail to encapsulate the effect of regime shifts, external shocks and real disturbances. MacDonald and Taylor (1994) employ a VAR analysis to data from 1976 to 1990 and discern three significant cointegrating vectors. However, only a single vector could not reject the imposed coefficient restrictions and was subsequently examined in a VECM framework that managed to beat the random walk both in-sample and out-sample.

occurs when stock/equity yields exceed money yields and induce a decline in the demand for money.

The relationship between PPP and cross-country productivity differentials was brought to the fore by Harrod in 1933 but was more notably examined as a construct of exchange rate equilibrium by Balassa and Samuelson in 1964. Balassa (1964) defines cross border productivity differences as differences in non-traded goods (services) and wages and subsequently shows that the cost of services in high income economies is greater than in low-income economies and that there is a positive correlation between Gross National Product (GNP) and the exchange rate. Similarly, Samuelson (1964) addresses currency overvaluation and the balance of payments by examining the indirect effect of production and wages on the exchange rate equilibrium. Samuelson (1964) deduces that an overvalued domestic currency impedes domestic employment, prevents a wage rise abroad and curbs private current account growth that subsequently hinders private investment as well as government spending. The role of productivity differentials in the determination of exchange rates, dubbed the Harrod-Balassa-Samuelson (HBS) effect, has been studied widely.[4]

Kim and Roubini (2000) use a restricted VAR to model the structure of the economy and the response of monetary bodies to monetary policy shocks by incorporating the global oil price to insulate 'exogenous' shifts in monetary policy. They show that the results of non-US monetary policy shocks on exchange rates as well as added macroeconomic variables

[4] Rogoff (1992) concretely tests for the HBS effect by constructing an intertemporal model of the real USD/JPY exchange rate. The model considers traded and non-traded goods output, government consumption expenditure and terms of trade shocks. Rogoff (1992) finds evidence of non-linear reversion with persistent shocks closer to exchange rate equilibrium. Obstfeld et. al (1996) enhance Rogoff's (1992) model to include a third aspect of production, trade costs, and corroborate the HBS effect subject to the factor-price equalization notion of trade theory.

support a wide variety of theoretical models by exhibiting stable money demand functions and upholding Purchasing Power Parity (PPP).[5]

In this chapter the scope of study regarding the exchange rate and monetary relationship is elaborated upon by proffering a hybrid model of the USD/CAD and the USD/GBP exchange rate that considers the failure of PPP and the HBS effect by integrating monetary as well as real features of the economy to overcome the failure of PPP and unstable money demand functions.

We commence our study, in section 2, by application of the conventional RID model (Frankel, 1979a) that defines money demand functions in terms of a ratio relative to domestic and foreign prices. The RID model is a justifiable specification in that it is able to identify fundamentals in the monetary sector. We compare results of the RID model with those of the hybrid model where money demand functions are based on real economic agents that have previously been branded as obstacles to PPP. The hybrid model introduces real stock prices in the money demand function as a proxy for real economic activity and since share prices respond to each other and because there is an underlying globalised economy driven by global fundamentals as well as global shocks to financial markets, we can capture characteristics of the long-term asset. The inclusion of a government spending ratio, productivity differential and real oil price are expected to shed more light on the determination of the long-run CAD/USD and GBP/USD exchange rates.[6]

[5] Bernanke et. al (1997) incorporate a global oil price because shocks to the oil price are not only most obvious but also anticipate USA recessions (recessions in the mid 70's, early 80's and early 90's). The outcome suggests that the impact of oil price shocks on the economy stems from restrictive economic policy rather than the behaviour of the oil price.
[6]By inclusions of a real oil price we expect to discern an unequivocal relationship between money and the CAD/USD exchange rate as Canada is a noteworthy producer of oil and 98% of its' energy is exported to the United States.

Our hybrid model for Canada - to differ from Backus (1986) - includes share prices, a productivity differential and a real oil price as directed by the array of current theory and applied studies. In light of Canada's long exposure to substantial and volatile capital flows in the era of flexible exchange rate regimes, we consider the CAD/USD exchange rate an ideal candidate for the monetary model of exchange rates as the majority of Canada's trade is carried out with the United States. We in addition consider that similarities in economic structure and economic expansion can minimize the amount of extraneous noise in the exchange rate model (Spiro, 1990). Considering economic growth and greater capital mobility one expects that the market for CAD is efficient and speculators hold enough capital to provide long-term arbitrage (Longworth, 1981).[7]

We in addition construct and modify the GBP/USD as well as the CAD/USD monetary models of the exchange rate for comparison with each other as the UK and Canada have a similar economic structure and growth rate but divergent trade policies with the United States. Our sample period is chosen in order to reflect on the impact of financial market deregulation in the late 1970's and early 80's in the USA, Canada and the UK and the effect of such a policy on stock prices and the exchange rate.

The rest of this chapter is ordered as follows. Section 2.2 reviews the theoretical monetary models of exchange rates and presents adapted specifications of the monetary model. Section 2.3 outlines the econometric approach to be applied in the study and describes the data. The empirical results are reported in Section 2.4 as well as the results of various tests that affirm the credibility of our approach. Finally, Section 2.5 provides conclusions and practical recommendations for further studies.

[7] Amano and Van Norden (1998) examine the impact of terms of trade on the CAD/USD exchange rate from 1973 to 1992 and find that terms of trade are long-run cointegrated with the real CAD/USD exchange rate and that causality runs from terms of trade to the real exchange rate.

2.2. The Monetary Models of Exchange Rate Determination

2.2.1. The Standard Monetary Models of Exchange Rates

The monetary approach to exchange rate modelling gathered momentum in the 1970’s and rests upon a stable money demand function and the validity of PPP (Sarno and Taylor, 2002; Taylor, 2007). The genre of classical models comprises the flexible price monetary model (Frenkel, 1976 and Bilson,1978), the sticky price model (Dornbusch, 1976; Hooper and Morton, 1982), the real interest differential monetary model (RID) (Frankel, 1979a), the modified flexible price monetary model (MFPM) (Frenkel, 1976; Mussa, 1976 and Bilson, 1978) and the modified real interest differential model (MRID) (Hunter and Ali, 2014). For the purposes of this study we examine versions of the RID and MRID model.

The flexible-price monetary model assumes flexible prices, instantaneous adjustment and the validity of PPP. Frenkel’s (1976) specification adopts a Cagan-type money demand function on real income and the interest rate in both the domestic and foreign market that lead to money market equilibrium (Frenkel and Koske, 2004):

$$m_t - p_t = \alpha_1 y_t - \alpha_2 i_t \tag{2.1}$$

$$m_t^* - p_t^* = \alpha_1 y_t^* - \alpha_2 i_t^* \tag{2.2}$$

where m_t is the money supply, p_t is the price level proxied by the consumer price index (CPI), i_t is the nominal interest rate, and y_t is real income. * denote the foreign country variable. Aside from the nominal interest rates all variables are expressed in terms of natural logarithms. We assume that the income elasticity α_1 and the interest rate semi-elasticity of money demand α_2 are indistinguishable at home and abroad. In terms of our study, we define the USA as the domestic market and both the UK and Canada as the foreign markets.

The PPP hypothesis if we assume that returns are continuously compounded implies that:

$$e_t^* = p_t - p_t^* \tag{2.3}$$

where e_t^* is the exchange rate (foreign currency per unit of domestic currency, GBP and CAD per USD in this instance). Eq. (2.3) suggests that testing the stationarity properties of the real exchange rate is analogous to testing for PPP.

Therefore, by repositioning Eqs. (2.1) and (2.2) we obtain equations for the domestic and foreign price levels and by further repositioning the equation for the relative prices levels we obtain:

$$p_t - p_t^* = (m_t - m_t^*) - \alpha_1(y_t - y_t^*) + \alpha_2(i_t - i_t^*) \tag{2.4}$$

By substitution of the equation for the spot exchange rate, Eq. (2.3), into the equation for relative prices of the domestic and foreign markets Eq. (2.4) we derive the flexible-price monetary model as follows:

$$e_t^* = \beta_1(m_t - m_t^*) - \beta_2(y_t - y_t^*) + \beta_3(i_t - i_t^*) + \varepsilon_t \tag{2.5}$$

Where ε_t is a white noise error term. α_1 and α_2 in Eq. (2.4.) are replaced by β_2 and β_3 in Eq. (2.5) since this is the reduced form of the model. Eq. (2.5) is the flexible-price monetary specification, developed by Frenkel (1976), Mussa (1976) and Bilson (1978), that describes the nominal exchange rate as a function of relative money supply, relative income and an interest differential. Therefore, a surge in domestic money supply corresponding to money supply abroad triggers a rise in domestic price levels and consequently a one-for-one depreciation of the nominal exchange rate. On the other hand, increased domestic output with respect to output abroad triggers a surge in the demand for money and consequently a reduction in the inflation rate prompting a re-valuation or appreciation of the exchange rate and a return to PPP. An increase in the domestic interest rate with respect to the foreign rate spurs a decline in real money demand that triggers an inflation rate rise and hence a currency depreciation once again prompting a return to PPP.

Eq. (2.5) implies that the coefficient on the relative real money supply is $\beta_1 = 1$, that the coefficient on relative real income is $\beta_2 > 0$ and that the coefficient on the relative interest rate is $\beta_3 > 0$.

The sticky-price monetary models assume that return to exchange rate equilibrium is slow so incorporates a term for expectations so that the nominal interest rate in Eq. (2.5) can be broken down into two parts namely the real interest rate r_t and the expected inflation rate Δp^e:

$$i_t = r_t + \Delta p^e \qquad (2.6)$$

$$i_t^* = r_t^* + \Delta p^{*e} \qquad (2.7)$$

Assuming interest rates are equal at home and abroad, it follows:

$$i_t - i_t^* = \Delta p^e - \Delta p^{*e} \qquad (2.8)$$

Substituting Eq. (2.8) into Eq. (2.5), the flexible-price monetary model is expressed as follows:

$$e_t^* = \beta_1(m_t - m_t^*) + \beta_2(y_t - y_t^*) + \beta_3(\Delta p^e - \Delta p^{*e}) + \varepsilon_t \qquad (2.9)$$

Frankel (1979a), recognised that exchange rates do not adjust instantaneously to long-run expectations and that the long-run equilibrium level is not only a function of long-run expectations but of short-term dynamics. Since expectations are unobservable they are interpreted as an inflation differential $(\Delta p^e - \Delta p^{*e})$ that considers the short-term dynamics of monetary policy. Thus, the expected change in the spot rate is expressed as a function of long-run expectations and short-run dynamics as follows:

$$E(\Delta e^*) = \delta(\bar{e}^* - e^*) + (\Delta p^* - \Delta p^{*e}) \qquad (2.10)$$

where δ is the speed of adjustment towards the exchange rate equilibrium level, Δp^e and Δp^{*e} express the domestic and foreign expected long-run inflation rates, respectively. Eq. (2.10), introduces the inflation rate differential that accounts for the short-run effect of monetary fundamentals on the exchange rate equilibrium level. The velocity of adjustment in the short-term, which depends on price stickiness, to a long-run exchange rate mean is equal to δ. Thus, it can be inferred that the difference between the long-run exchange rate mean and the spot rate $(\bar{e}^* - e^*)$ are comparable to the difference between expected inflation rates at home and abroad $(\Delta p^e - \Delta p^{*e})$.

Introduction of uncovered interest parity (UIP) $E(\Delta e^*) = i_t - i_t^*$ postulates that information on expectations is efficient. Thus, investors do not require compensation for uncertainty and returns are expected to equalize across borders deducing that assets are perfect substitutes. By merging the assumption of UIP with Eq. (2.10) and repositioning we reach an exchange rate equation as follows:

$$e^* = \bar{e}^* - \frac{1}{\delta}[(i - \Delta p^e) - (i^* - \Delta p^{*e})] \qquad (2.11)$$

The long run exchange rate ē is procured from the flexible-price monetary model in Eq. (2.9) that in a quasi-reduced form can be expressed as $e_t^* = (m_t - m_t^*) + \alpha_1(y_t - y_t^*) + \alpha_2(\Delta p^e - \Delta p^{*e}) + \varepsilon_t$ (Macdonald, 2007) so by substitution of ē in Eq. (2.11) we obtain:

$$e^* = (m - m^*) + \alpha_1(y - y^*) + \alpha_2(\Delta p^e - \Delta p^{*e}) - \frac{1}{\delta}[(i - \Delta p^e) - (i^* - \Delta p^{*e})]. \qquad (2.12)$$

Traditionally a proxy for the real interest rate is a short-term interest rate (e.g. the discount rate) that captures short-run monetary dynamics such as liquidity shocks. A long-term bond rate is employed as a proxy for long-run expectations about inflation rates (Frankel, 1979a and Macdonald, 2007). The quasi-reduced specification of the standard real interest differential model is:

$$e_t^* = \beta_1(m_t - m_t^*) + \beta_2(y_t - y_t^{s*}) + \beta_3(i_t^* - i_t^{s*}) + \beta_4(i_t^* - i_t^{l*}) + \varepsilon_t \qquad (2.13)$$

where i_t^s represents the short-term interest rate and i_t^l the long-term interest rate (e.g. 10-year government bond rate). Here a star (*) denotes the foreign country, in our model Canada and the UK (while the USA is home). The RID model in Eq. (2.13) posits that the exchange rate is conditioned on foreign and domestic money supply so that a money supply surge in the domestic economy relative to that abroad would raise the domestic inflation rate consequently triggering a one-for-one depreciation (β_1=1) of the domestic currency relative to that abroad. Growth in domestic output as well as a drop in the anticipated rate of domestic inflation creates a surge in the domestic demand for money thereby triggering an appreciation (β_2>0 and β_4<0) of the home currency relative to the foreign currency. A rise in the domestic interest rate with respect to the foreign rate would in addition prompt a surge in money supply and hence an appreciation (β_3>0) of the home currency with respect to the foreign currency. Thus, the model infers that when the exchange rate is denoted in terms of the foreign country the coefficient on relative money is equal to unity (β_1=1), the coefficients on income (β_2>0) and the short interest rate (β_3>0) are positive while the coefficient on the long-interest rate is negative (β_4<0) (Frankel, 1979a).

2.2.2. Variants of The Monetary Models of Exchange Rates

Since the conventional RID model assumes that adjustment to equilibrium is instantaneous, only variables that account for short run dynamics are included. Friedman

(1988) recognized the shortfall of the classical model after exhibiting the inverse relationship between the Dow Jones stock market index and the velocity of a USA monetary aggregate, sum as M2. Many others have followed in Friedman's (1988) path and by adding stock prices to the money demand equations have tightened the gap between short-run dynamics and long-run expectations and enhanced model stability. Solnik (1987) examines the relationship between exchange rates, monetary policy and economic activity in a study comprising the USA, UK, Canada, France, Switzerland, Germany, the Netherlands and Japan as those countries represented 90% of world market capitalization. Solnik (1987) found a weak relationship between stock price differentials and exchange rate variance but concluded that stock prices encompass views about the future and perceived economic growth would have a positive influence on the exchange rate. McCornac (1991) showed in a study for Japan, between 1975 and 1987, a correlation between lagged equity variables and money demand in addition to a risk spreading effect whereby an increase in the procurement of risky assets is accompanied by a demand for secure short-term assets in addition to money. Among others who have affirmed that stock prices enhance the stability of money demand functions are: Choudhry (1996) for the CAD/USD in the post WW2 period until 1989 finds that stock prices exhibit a marked effect in the determination of stationary real M1 and M2 demand functions in the long-term in the U.S.A as well as Canada; Thornton (1998) for Germany from 1960 to 1989 who found firstly, that share prices exhibit a positive wealth effect on the long-run demand for money secondly, feedback effects amongst real money balances and interest rates and thirdly, unidirectional Granger causality from real income to interest rates, from interest rates to stock prices and from money to income; Caruso (2001) for a panel of 19 industrial countries and 6 developing countries who found a negative wealth effect for the UK, Switzerland and Japan as well as a substitution effect in

Italy and to a lesser extent in France and Germany; Baharumshah et al. (2002) for the RM vis-à-vis the USD and JPY between 1983 and 1996 show that a restricted VAR model with imposed restrictions of exogeneity, demonstrates cointegration, parameter stability and adjustment of the exchange rate to disequilibrium in the long-run relation; Morley (2007) for the UK between 1984 and 2002 finds that in the short-run stock prices influence the exchange rate and that in equilibrium the monetary model is stable in the long-run; Baharumshah et al. (2009) for China between 1990 and 2005 shows a long-term relation amongst M2 and income, inflation, foreign interest rates and stock prices that have a wealth effect on long-run and short-run broad money demand.

A more recent contribution to the literature on the relationship between exchange rates and share prices has been made by Hunter and Ali (2014) who enhance a conventional RID monetary model of the JPY/USD exchange rate between 1980 and 2009. Hunter and Ali (2014) attribute improved model performance to the addition of share prices, government spending, a productivity differential and a global oil price that account for real economic activity as well as long-run expectations and yield a unique cointegrating vector.

To reiterate the relationship between money demand and stock prices is constituted by a wealth effect and a substitution effect. The wealth effect describes an increase in money demand because of increased nominal wealth, a rise in the number of financial market transactions and surging share prices that indicate greater risk appetite counteracted by short-term hedges in money or short-term protected assets. On the other hand, the substitution effect implies that higher share prices yield higher returns than holding money so that money is replaced by shares.

In our model aside from consideration of long-run expectations we include stock prices in light of literature that demonstrates a distinct relationship between exchange rates and stock prices. Some of this literature is discussed by Aggarwal (1981) who finds that exchange rate fluctuation is dependent upon whether the domestic country is predominantly an exporter or an importer since an appreciation of the home currency would discourage exporters and increase import of production input so that the effect of a fluctuation will differ depending on the size of a producers' assets and liabilities. An appreciation of the home currency for a country that mainly exports would induce a decline in stock prices and vice versa and by so doing decrease the price of domestic stock, the opposite is true for a currency depreciation.

Jorion (1990) suggests that an inverse relationship between exchange rates and stock prices may exist and that a currency appreciation would prompt a fall in stock prices in view of diminished returns, contingent upon exchange rate exposure. Bahmani-Oskooee and Sohrabia (1992) in an examination of the relationship between the effective USD exchange rate and the S&P 500 stock index between July 1973 and December 1988 observe a bidirectional relationship between the exchange rate and stock index but no evidence of long-run cointegration. Bodnar and Gentry (1993) look at the relationship between industry and exchange rate behaviour in the USA, UK and Canada and find that exchange rate exposure is determined by industry returns and therefore stock prices.

It is in addition observed that Canadian and Japanese industries are more affected by exchange rate fluctuation than those in the USA, most likely since they are smaller and more inclined internationally. Bartov and Bodnar (1994) fail to provide evidence of a causal relationship between atypical stock returns and contemporaneous USD exchange rate variance but, find a negative relationship between lagged exchange rate variance and atypical stock returns owing to late publication of firm's assets and liabilities. Granger et al. (2000) apply unit root tests and a cointegration technique to a set of East Asian countries between 1986 and 1998 where exchange rates are shown to guide share prices in South Korea and the opposite in the Philippines; however, evidence from Indonesia and Japan is inconclusive.

Nieh and Lee (2002) investigate daily time series between 1993 and 1996 for Canada, France, Germany, Italy, Japan, UK and the USA by employing the Johansen (1995) and Engle-Granger (1987) two step methodologies and most interestingly find no causal relationship between stock prices and exchange rates. Caporale et al. (2002) test for causality between stock prices and the Indonesian rupiah (IDR), JPY, South Korean won (KRW) and Thai baht (THB) vis-à-vis the USD using a multivariate GARCH model specified by Baba et al. (1990). They find in the sample period, covering the 1997 East Asian crisis, a negative relationship in the JPY and KRW as compared to a positive relationship in the IDR and THB with bilateral effects being experienced in the IDR and THB following the 1997 crisis.

Phylaktis and Ravazzola (2005) examine the relationship between relative stock prices and the exchange rate in Hong Kong, Malaysia, Singapore, Thailand and the Philippines, by employing cointegration techniques and a Granger causality test. They find a positive relationship between stock prices and exchange rates that is driven by the USA stock market. Caporale et al. (2014) survey 6 industrialized countries in periods preceding and succeeding the 2007-2010 financial crisis and report a unidirectional relationship between stock prices and exchange rates for the USA and the UK where exchange rates are led by stock prices but an inverse relationship for Canada. Caporale et al. (2014) in addition report that diversification for domestic industries would be an ineffective tool for hedging during financial turmoil in view of the tie between domestic production, stock prices and the exchange rate.

The relationship between exchange rates and stock prices as discussed above is demonstrated for the GBP vis-à-vis the USD and the CAD vis-à-vis the USD, respectively, in Figures 2.1. and 2.2. below. Figures 2.1. and 2.2. display quarterly relative real stock prices for the UK and Canada deflated by corresponding CPIs as well as the variance of the GBP/USD and CAD/USD exchange rates over the period 1980:Q1-2015:Q4. Observation of the graphs indicates that shifts in the relative real stock price may indicate behavioural patterns for the USD against the GBP and CAD.

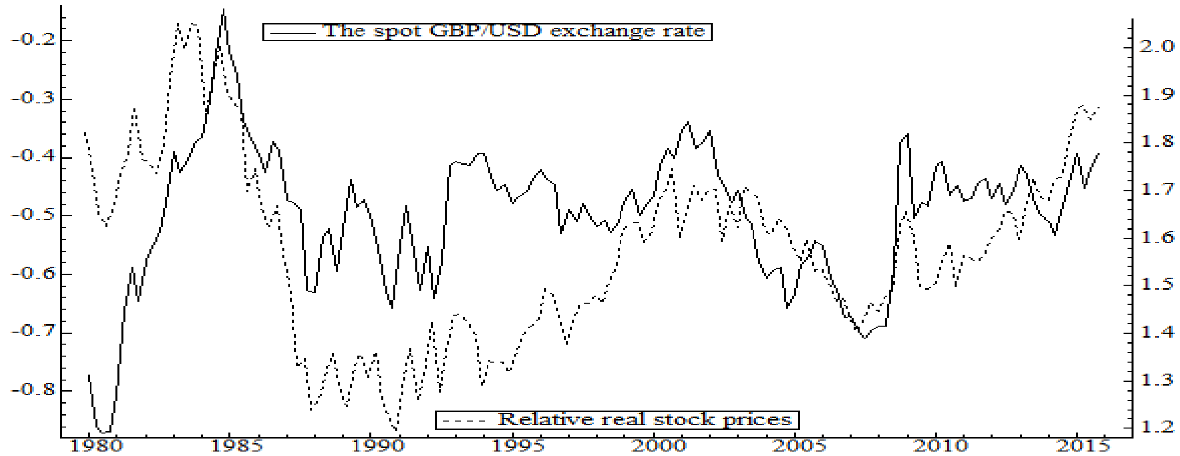

Figure 2.1. The Movements of the Spot GBP/USD Exchange Rate (Left-Hand Scale) and Relative Real Stock Prices (Right-Hand Scale) Between the USA and the UK From 1980:Q1 to 2015:Q4.

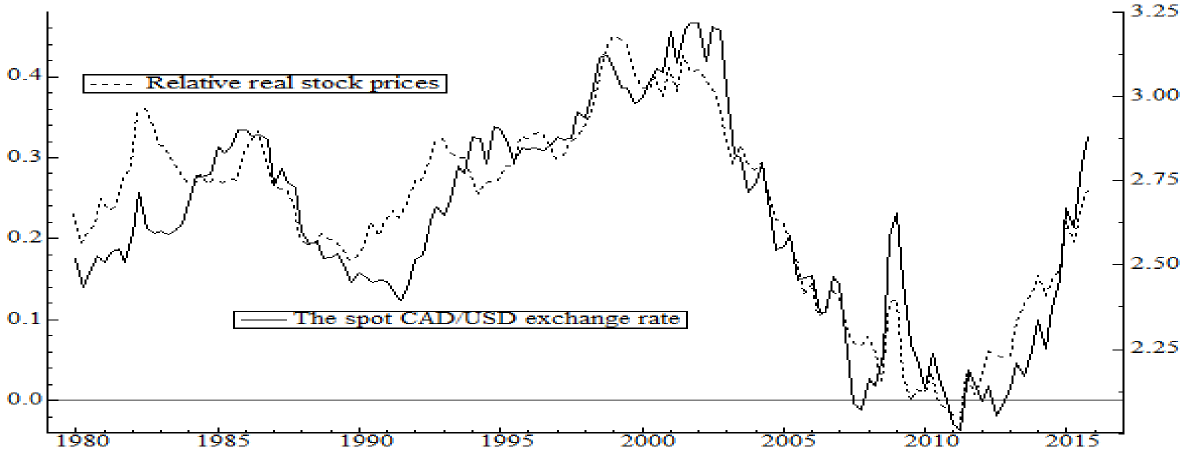

Figure 2.2. The Movements of the Spot CAD/USD Exchange Rate (Left-Hand Scale) and Relative Real Stock Prices (Right-Hand Scale) Between the USA and the Canada from 1980:Q1 to 2015:Q4.

The graphs indicate an apparent negative correlation between relative real stock prices and the GBP/USD and CAD/USD exchange rates during periods of crises specifically in the early 1980's recession, the 1987 stock market collapse, the dot-com crash in the early 2000s and the global financial crisis of 2007-2008. A negative wealth effect (Friedman, 1988) is evident as relative real stock prices rise as the GBP and CAD depreciate against the USD. The dominance of domestic industry determines the effect on productivity so that an appreciation of the domestic currency, in this instance the USA, for an economy that primarily exports would prompt a fall in stock prices and vice versa.

Founded on evidence documenting the tie between stock prices and exchange rates the money demand equations (2.1) and (2.2) are augmented as follows:

$$m_t - p_t = \alpha_1 y_t - \alpha_2 i_t + \alpha_3 s_t, \qquad (2.1)'$$

$$m_t^* - p_t^* = \alpha_1 y_t^* - \alpha_2 i_t^* + \alpha_3 s_t^*, \qquad (2.2)'$$

Where s_t and s_t^* correspondingly denote the domestic and foreign real stock price indices in logarithmic form. The coefficient sign on the real stock price, α_3, is contingent upon whether a wealth effect or substitution effect directs the money demand function.

Following on from our specification of the money demand function and discussion regarding its long-run stability we continue building our monetary model with a specification for PPP and discuss factors affecting its validity. For PPP to hold it is a requirement that the real exchange rate under consideration be stationary. Stationarity properties are generally examined by implementation of unit root tests (Taylor, 1988 and Mark, 1990), variance ratio tests (Huizinga, 1987), fractional integration methods (Diebold et al., 1991 and Cheung and Lai, 1993b) and tests of cointegration between the real exchange rate and national price levels.

In the current age of flexible exchange rate regimes, the Johansen (1995) test for cointegration and Engle and Granger (1987) two-step methodology have generally enjoyed little success in evidencing a meaningful link between real exchange rates and relative national price levels. Implementation in a multivariate framework (Kugler and Lenz, 1993) has yielded some evidence of a long-run relationship between exchange rates and relative price levels but given the bias of measurement error in prices as well as proportionality and symmetry conditions (Cheung and Lai, 1993b) many results are deemed inconclusive. Fisher and Park (1991) examine the relationship between consumer as well as industrial price indices and exchange rates for 11 major economies from March 1973:M1 to May 1988:M12. They show by means of cointegration analysis a link between prices and the exchange rates for all economies in the study aside from the USA and Canada. The estimated error correction models show that equity prices as compared to trade prices drive exchange rate adjustment and support the idea that relative price levels are determined by unrecognizable economic activity such as long-run expectations. Coakley et al. (2005b) study PPP by application of Pesaran's (2007) cross-sectional enhanced Im et al. (1997) test that tolerates cross sectional dependence (CSD) to a panel of 15 OECD countries and perceives that CSD is not key to the failure of PPP but rather selection of price index whether that be the CPI, the production price index (PPI) or the industrial production index (IPI).

Edison et al. (1997) implement a Johansen (1995) methodology employing Monte Carlo experiments to a set of data from 14 industrialized countries from 1974:Q1 to 1992:Q4 and exhibit evidence of cointegration that is rejected based on proportionality and symmetry conditions. Hall et al. (2013) test for homogeneity in prices based on the assumption that the gap between exchange rates and prices is driven by specific factors and if those factors were to be neutralized we would expect exchange rates and prices to act in proportion to one another. Hall et al. (2013) by application of a time-varying coefficient methodology to 9 Euro area countries as well as Japan, Mexico and Canada find that price coefficient averages are homogeneous in the long-term and therefore conclude in absolute favour of PPP. Bahmani-Oskooee et al. (2015) develop a panel test for stationarity that considers acute structural breaks as well as smooth transitions and apply it to a panel of 34 OECD countries. They find support for PPP amongst 17 of the 34 countries. Similarly, Wang et al. (2016) apply a panel test for stationarity that considers cross sectional heterogeneity to 20 developing and 20 industrialized nations and test for the Harrod-Balassa-Samuelson (HBS) effect finding evidence of PPP in the panel of developed countries where productivity growth triggers a currency appreciation. Donayre and Panovska (2016) apply a Bayesian threshold vector autoregression to examine the link between exchange rate pass-through and economic activity in Canada and Mexico observing that pass-through is reliant upon the state of the economy and nonlinearities exist so that the greater the pass-through coefficient the greater output growth.

Conventionally, unit root tests have yielded non-stationary exchange rates that assume a random walk (Taylor, 1988; Mark, 1990; Sarno and Taylor, 2002 and Taylor, 2007). A case of particular note pertains to the USD/JPY exchange rate that has more often than not proved to be non-stationary and hence construction of a well-grounded monetary model elusive (Macdonald and Nagayasu, 1998; Rogoff, 2001; Caporale and Pittis, 2001; Lizardo and Mollick, 2010). However, Chortareas and Kapetanios (2004) find that by application of a non-linear type Augmented Dickey-Fuller test founded on an application of an ESTAR model that the USD/JPY is stationary because the real JPY exchange rate vis-à-vis more G7 (group of the seven most advanced economies in the world including America, Canada, Japan, Great Britain, Germany, France and Italy) and Asian currencies demonstrated mean reversion in the period following the collapse of Bretton Woods. Unit root test results have shown improvement when applied to data spanning over an extended length of time (Frankel, 1986; Edison, 1987; Abuaf and Jorian, 1990; Lothian and Taylor, 1997 and 2000; Sarno and Taylor, 2002) and upon application of a panel data methodology (Flood and Taylor, 1996; Frankel and Rose, 1996; Wu, 1996; Chinn and Johnston, 1997; Lothian, 1997; Papell, 1997; Mark and Sul, 2001; Rapach and Wohar; 2004; Coakley et al., 2005a; Jiang and Bahmani-Oskooee, 2015 and Wang et al., 2016)

The breakdown of PPP has been studied extensively and ascribed to the impact of real economic activity. Balassa (1964) and Samuelson (1964) pioneered the notion that the difference between traded and non-traded goods could have an equal if not greater impact on the relationship between real exchange rates and price levels. Samuelson (1964) posits that a rise in production of traded goods is likely to raise the cost of labour in the sector as well as in the non-traded goods sector since labourers are interchangeable and free to move. Hence, we perceive a relationship between non-traded goods (services), government spending and the real exchange rate. The real oil price has also been incorporated in preceding monetary models (Rogoff, 1992; Chinn and Johnston, 1997; Bernanke et al., 1997; Kim and Roubini, 2000) to account for terms of trade shocks and insulate 'exogenous' shocks to monetary policy.

The inadequacy of PPP to substantiate behaviour of exchange rates as discussed above is demonstrated in Figures 2.3. and 2.4. below. Figures 2.3. and 2.4., correspondingly, show that the real GBP/USD and CAD/USD exchange rates are persistent as well as non-mean reverting and do not exhibit similar trends to the national price ratio and thus national prices do not predict exchange rate movements. The breakdown of PPP is shown clearly by the failure of relative national prices to powerfully explain the spot GBP/USD and CAD/USD exchange rates.

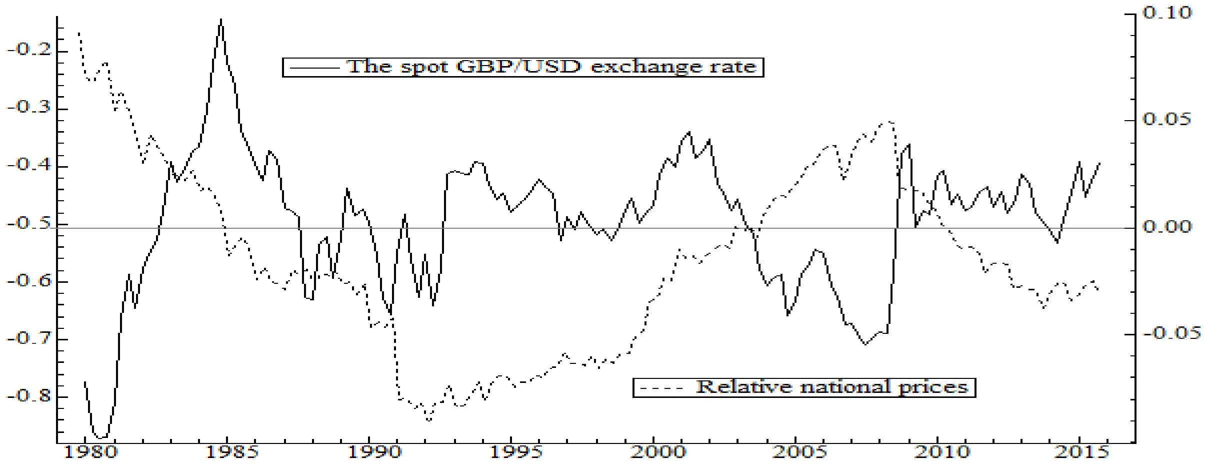

Figure 2.3. The Behaviour of the Spot GBP/USD (Left-Hand Scale) Exchange Rate and Relative National Prices (Right-Hand Scale)

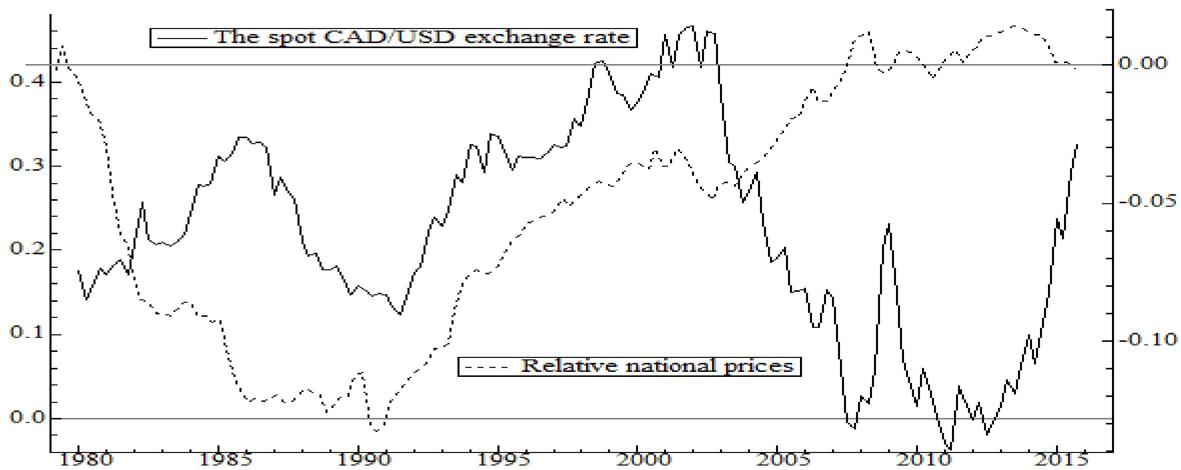

Figure 2.4. The Behaviour of the Spot CAD/USD (Left-Hand Scale) Exchange Rate and Relative National Prices (Right-Hand Scale)

Rogoff (1992) concluded that his structural model of the JPY/USD failed to capture effects of external shocks and real disturbances. We expand upon Rogoff's (1992) study and account for the effect of external shocks and real disturbances on the persistence of the GBP/USD and CAD/USD exchange rates by incorporation of stock index prices, a productivity differential, a government spending ratio and the effect of a real oil price. Together the variables should account for short and long-term dynamics of the economy as well as terms of trade.

Considering the HBS effect foreign and domestic price levels are broken down into prices of traded p_t^T and non-traded goods p_t^{NT} :

$$p_t = (1-\alpha)p_t^T + \alpha p_t^{NT} = p_t^T + \alpha(p_t^{NT} - p_t^T) \tag{2.14}$$

$$p_t^* = (1-\alpha)p_t^{*T} + \alpha p_t^{*T} = p_t^{*T} + \alpha(p_t^{*NT} - p_t^{*T}) \tag{2.15}$$

Where, α(1−α) represents the amount of non-traded (traded) goods in the economy. The real exchange rate is balancing the nominal exchange rate with the national price ratio for the domestic and foreign economy:

$$q_t = e_t^* - p_t + p_t^* \tag{2.16}$$

Where q_t represents the real exchange rate. By means of substituting the traded and non-traded goods price ratios from Eq. (2.14) and Eq. (2.15) into Eq. (2.16), the real exchange rate is presented in Eq. (2.17) as follows:

$$\begin{aligned} q_t &= e_t^* - [p_t^T + \alpha(p_t^{NT} - p_t^T)] + [p_t^{*T} + \alpha p_t^{*NT} - p_t^{*T})] = \\ &(e_t^* - p_t^T + p_t^{*T}) - \alpha[(p_p^{NT} - p_t^T) - (p_t^{*NT} - p_t^{*T})] \end{aligned} \tag{2.17}$$

Since arbitrage only pertains to traded goods it follows that PPP is pertinent to the traded goods sector. Therefore $(e_t^* - p_t^T + p_t^{*T})$ in Eq. (2.17) should equal zero. The real exchange rate is represented in relation to traded and non-traded goods in Eq. (2.18):

$$q_t = -\alpha[(p_t^{NT} - p_t^T) - (p_t^{*NT} - p_t^{*T})] \tag{2.18}$$

We observe from Eq. (2.18) that the real exchange may be derived from the comparative prices of traded and non-traded goods in the domestic and foreign market. Positive movements in the home prices of traded and non-traded goods compared with prices abroad brings about an appreciation of the real exchange rate. Engel and Rogers (1994) found that variation in the price of traded and non-traded goods across borders accounts for much of the failure of the Law-of-One-Price in an examination of CPI data for USA and Canadian cities. Rogers and Jenkins (1995) find support for PPP in the traded goods sector, so that real exchange rate persistence stems from relative price fluctuations within countries but ascribe the breakdown of PPP in Canada and the United States to the presence of non-traded goods in the general price index. Parsley and Wei (1996) in a panel of 51 prices from 48 US cities as well as Crucini and Shintani (2008) in a study of prices from major cities in 63 countries including the US, observed that reversion to parity is homogeneous and is faster for traded goods and larger price differentials. As compared to Parsley and Wei (1996), Imbs et al. (2005) attribute the failure of PPP to heterogeneous dynamics and find that the persistence of disaggregated relative prices is generally smaller than the persistence of the real exchange rate. While Parsley and Wei (1996) observe that the speed of convergence towards PPP is much faster in the traded goods sector as compared to the non-traded goods sector. Similarly, Cecchetti et al. (2002) examine data from 19 U.S. cities in a panel over the period 1918-1995 and observe a slow rate of reversion to parity, which is ascribed to variances in the rate of adjustment to shocks of varying dimension as well as the incorporation of non-traded goods prices in the price index.

Strauss (1999) observes that increases in the prices of domestic non-tradeable products are correlated with appreciations of the real and nominal exchange rates and may substantiate long-term but not lasting deviations from PPP. Following Strauss (1999) who assumes a competitive market where industries define prices in consideration of labour costs:

$$p_t^T = w_t - prod_t^T \, , p_t^{NT} = w_t - prod_t^{NT} \, , p_t^{NT} = w_t^* - prod_t^{*T} \, , p_t^{*NT} = w_t^* - prod_t^{*NT} \qquad (2.19)$$

where, w_t is the wage rate that is identical in the traded and non-traded sectors since labour is interchangeable whereas, $prod_t^T$ and $prod_t^{NT}$ represent productivity in the traded and non-traded segments, correspondingly. We therefore determine that the relative price movements of nontraded goods are substantiated by the difference between productivity in the traded and non-traded goods' industries:

$$p_t^{NT} - p_t^T = prod_t^T - prod_t^{NT}, p_t^{*NT} - p_T^{*T} = prod_t^{*T} - prod_t^{*NT} \quad (2.20)$$

Substituting Eq. (2.20) into Eq. (2.18):

$$q_t = -\alpha[(prod_t^T - prod_t^{*T}) - (prod_t^{NT} - prod_t^{*NT})] \quad (2.21)$$

We observe from Eq. (2.21) that a surge in productivity in the traded sector relative to productivity in the non-traded sector prompts a decline in the prices of traded goods and consequently an exchange rate appreciation. Chinn (1997a) examining a panel of 14 OECD countries finds support for a monetary model of the real exchange rate that incorporates a measure for sectoral productivity that captures the long-run relation as well as a measure for government spending that captures short-run dynamics. This happens because government spending shocks to the prices of traded and non-traded goods will exert a greater impact on the exchange rate than shocks emanating from the private sector. Government expenditure is predominantly allocated to non-traded goods so that a surge in government spending would prompt a rise in the prices of non-traded products and a subsequent currency appreciation. Thus, Eq. (2.21) is modified accordingly:

$$q_t = -\alpha[(prod_t^T - prod_t^{*T}) - (prod_t^{NT} - prod_t^{*NT})] + \mu(gs_t - gs_t^*) \quad (2.22)$$

where, $gs_t(gs_t^*)$ represents domestic (foreign) government consumption expenditure as a percentage of gross domestic product (we anticipate μ to be negative, μ<0).

Rogoff (1992), Chinn and Johnston (1997), Bernanke et. al (1997), Kim and Roubini (2000) include a global oil price in their models of the exchange rate to account for terms of trade shocks. Bernanke et. al (1997) construct a monetary model of the USD exchange rate incorporating an oil price since the USA is a major producer of oil and shocks to the oil price influence the exchange rate in an obvious way and direct USA business cycles. The USA exports domestically produced oil when the global price of oil is high and utilizes domestically produced oil when profit from oil exports is negligible thereby influencing the value of the USA currency (Lizardo and Mollick, 2010). Other studies that observe the effect of the oil price on currency values include Lastrapes (1992), DeGregorio and Wolf (1994), Enders and Lee (1997), Deloach (2001), Chen and Chen (2007) and Zhang et al. (2008). We therefore modify Eq. (2.22) to include a global oil price:

$$q_t = -\alpha[(prod_t^T - prod_t^{*T}) - (prod_t^{NT} - prod_t^{*NT})] + \mu(gs_t - gs_t^*) + \gamma roil_t \qquad (2.23)$$

where, $roil_t$ is the real oil price. We employ the West Texas Intermediate (WTI) spot price and account for inflation by deflating the price by the rate of change of the USA Consumer Price Index (CPI). We anticipate ɣ to be negative ɣ < 0. Since quarterly data on non-tradable goods and services are difficult to gather we assume that $prod_t^{NT} = prod_t^{*NT}$. Chinn (2000), Lee et al. (2002), Miyakoshi (2003), Wang and Dunne (2003) and Tsen (2011) amongst other scholars show that productivity differentials account for real economic fundamentals that exert an impact on the nominal exchange rate. Therefore, by substitution in (2.23) we derive:

$$q_t = -\alpha(prod_t^T - prod_t^{*T}) + \mu(gs_t - gs_t^*) + \gamma roil_t \qquad (2.24)$$

An augmented flexible price model (FPM) is derived by substitution of Eq. (2.24) and the money demand functions represented in Eq. (2.1)’ and Eq. (2.2’) as follows:

$$e_t^* = \beta_1(m_t - m_t^*) + \beta_2(y_t - y_t^*) + \beta_3(i_t^s - i_t^{s*}) + \beta_4(s_t - s_t^*) + \beta_5(prod_t^T - prod_t^{T*}) + \beta_6(gs_t - gs_t^*) + \beta_7\gamma roil + \varepsilon_t \quad (2.5)'$$

The hybrid model is derived by further enhancing the augmented FPM in Eq. (2.5)' by incorporation of a long-term interest rate that accounts for long-term dynamics in the real economy. The hybrid model is presented in Eq. (2.13)' as follows:

$$e_t^* = \beta_1(m_t - m_t^*) + \beta_2(y_t - y_t^*) + \beta_3(i_t^s - i_t^{s*}) + \beta_4(i_t^l - i_t^{l*}) + \beta_5(s_t - s_t^*) + \beta_6(prod_t^T - prod_t^{T*}) + \beta_7(gs_t - gs_t^*) + \beta_8\gamma roil + \varepsilon_t \quad (2.13)'$$

where $(i^l - i^{l*})$ is the long-term interest rate differential.

When the exchange rate is denoted in terms of the foreign country, in our model an * denotes the foreign country, we assume that the coefficient on the relative money supply is equal to unity ($\beta_1 = 1$), the coefficient on relative income to be positive ($\beta_2 > 0$) and the coefficient on the relative short-run interest rate to be positive ($\beta_3 > 0$). By reference to the relative share price a wealth effect (capital flight) domestically would give rise to an appreciation of the USD so that the coefficient the relative share price would be positive ($\beta_4 > 0$) and a substitution effect would give rise to a depreciation of the home currency so that the coefficient on the relative share price would be negative ($\beta_4 < 0$). Balassa (1964) and Samuelson (1964) assume that a rise in the productivity differential of traded goods as opposed to non-traded products domestically lowers prices of traded goods at home and brings about an appreciation of the domestic currency. We therefore expect the coefficient on the productivity differential to be positive ($\beta_5 > 0$) when productivity in the domestic traded goods sector exceeds that in the non-traded sector, as here currency is in terms of the foreign country. Chinn (2000) as well as Kim and Roubini (2008) state that a rationale for considering government consumption in a monetary model of the exchange rate is that government expenditure is greatest in the non-traded good sector. We therefore assume that an increase in relative government spending would raise the price of nontradables and prompt an appreciation of the domestic currency so anticipate the coefficient on relative government expenditure to be positive ($\beta_6 > 0$) since in our model the currency is denoted in terms of the foreign country.

Amano and van Norden (1998) note that the oil price has a persistent effect on the USD. Lizardo and Mollick (2010) observe that a rise in the real oil price prompts a depreciation of the USD against the currencies of Canada, Mexico and Russia, all major oil exporting countries and vice versa against the currencies of Japan and Denmark, both oil importers. Canada is the world's fourth largest oil exporter [8] and most of its' exports are to the United States so we expect that this relationship should affect the persistence of the CAD/USD exchange rate. Global oil prices are quoted in terms of the USD so we expect that a rise in the real oil price would induce an appreciation of the USA currency, so that the coefficient on the real oil price would be positive ($\beta_7 > 0$). The variable coefficients are assumed to be of the opposite sign when the exchange rate is denoted in terms of the foreign currency.

To reiterate we anticipate that for a model of the exchange rate, currency denoted in terms of the foreign country, the coefficients are assumed to be $\beta_1 = 1$, $\beta_2 > 0$, $\beta_3 > 0$, $\beta_4 < 0$, $\beta_5 < 0$ (substitution effect) or $\beta_5 > 0$ (wealth effect), $\beta_6 > 0$, $\beta_7 > 0$ and $\beta_8 > 0$. The sign on the relative long-run interest rate coefficient and relative stock price coefficients depends on whether we observe a substitution effect or wealth effect. In section 2.4 we test the proficiency of the standard RID model in Eq (2.5)' as well as that of the hybrid model in Eq. (2.13)' in substantiating persistence of the nominal GBP/USD and CAD/USD exchange rates.

[8] http://www.nrcan.gc.ca/energy/oil-sands/18078

2.3. The Data and Econometric Approach

2.3.1. Data

The study draws on quarterly data that is, for most series in the sample, seasonally unadjusted for the United States vis-à-vis the UK and Canada over the period 1980:Q1 – 2015:Q4. Our analysis spans the period 1980:Q1 – 2015:Q4 to consider the implications of financial market deregulation following the fall of Bretton Woods and to gain greater depth of understanding of the factors that drive exchange rates in a flexible exchange rate regime. This sample period is marked by high capital mobility and integrated markets that facilitate cross-country borrowing and lending as well as smooth consumption throughout the business cycle (Obstfeld and Rogoff, 1995). The sample period related to the USA, Canada and the UK is in addition related to historically low interest rates following the global financial crisis of 2007-2008. Quarterly intervals are applied since Gross Domestic Product (GDP) data is not accessible monthly and using an alternative estimate for national income would compromise the model.

The exchange rates are the official rates denoted as CAD per USD and GBP per USD. The short-term interest rates are represented by money market rates while the long-term interest rates are 10-year government bond yields. A respective consumer price index (CPI) is used to deflate stock price indices in the USA, UK and Canada correspondingly represented by the S&P 500 index, the FTSE 100 and the S&P/TSX Composite Index. Government spending is defined as final government consumption expenditure as a percentage of gross domestic product while productivity is defined as the industrial production index divided by the corresponding employment rate. The real oil price is the West Texas Intermediate (WTI) Cushing crude oil spot price, dollars per barrel, deflated by the USA consumer price index. The exchange rates, interest rates, industrial production index and consumer price indices (CPI) are extracted from the International Monetary Fund's (IMF) International Financial Statistics (IFS), whereas money supply (M1) for the US and Canada and stock prices are collected from Thomson DataStream. Gross domestic product and final government consumption figures for the USA and UK in addition to employment rates are obtained from the Organisation for Economic Co-operation (OECD) main economic indicators (MEI) database. The West Texas Intermediate (WTI) Cushing crude oil spot price, USDs per barrel, is retrieved from Archival Federal Reserve Economic Data.

Gross domestic product and final government consumption expenditure figures are expenditure based, at market prices and volume estimates chained to the OECD base year 2010. Exchange rates, long-run interest rates, stock prices and the oil price are end of period data. All variables have been expressed in their logarithmic form except the interest rate differentials and the government spending ratio that is a percentage.

Graphs of the variables in levels (upper graphs) and first differences (lower graphs) are displayed below in Figures 2.5. - 2.12. for the Canadian variables and Figures 2.13. - 2.20. for the UK variables. The real oil price employed in both the GBP/USD and CAD/USD model is shown in Figure 2.21. below.

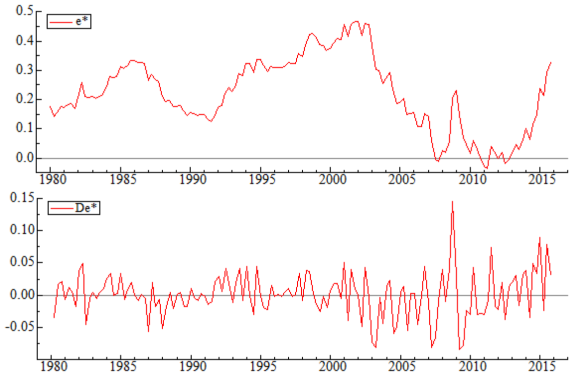

Figure 2.5. Spot CAD/USD Exchange Rate

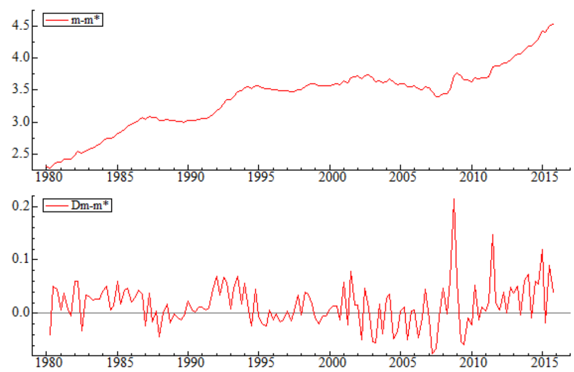

Figure 2.6. Relative M1 for the CAD/USD Model

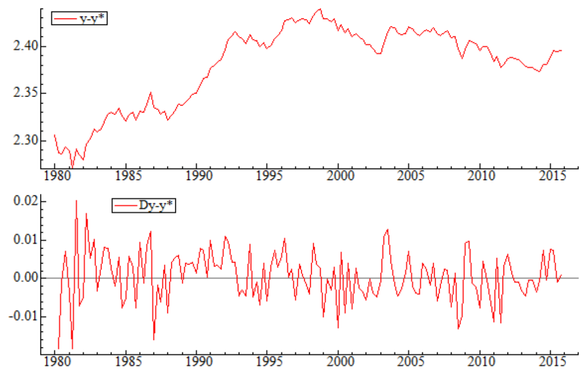

Figure 2.7. Relative GDP for the CAD/USD Model

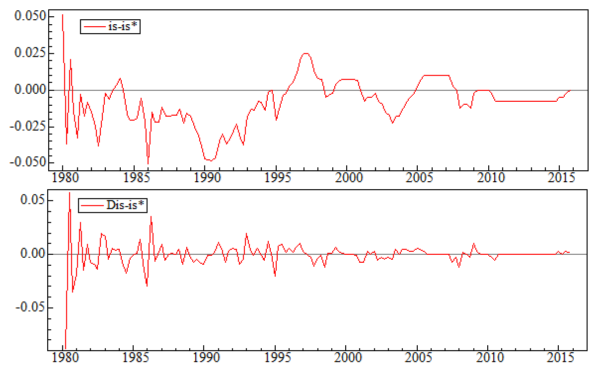

Figure 2.8. Relative Short-Term Interest Rate

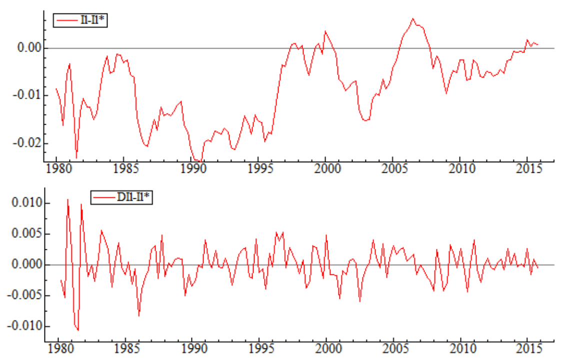

Figure 2.9. Relative Long-Term Interest Rate

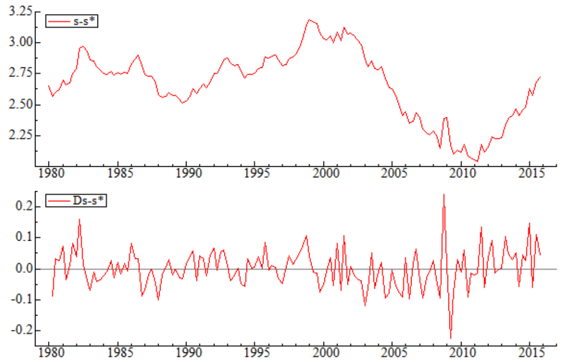

Figure 2.10. Relative Real Stock Prices

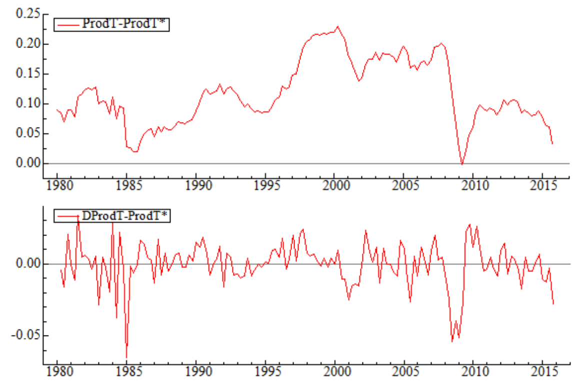

Figure 2.11. Productivity Differential

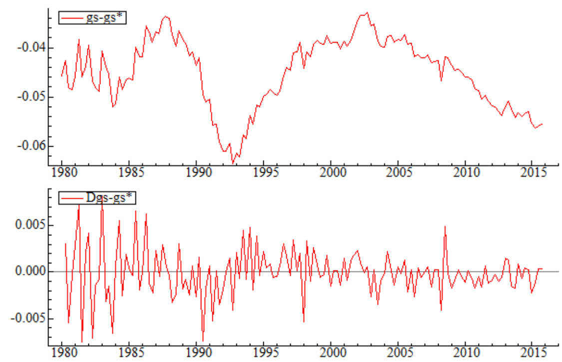

Figure 2.12. Government Spending Differential

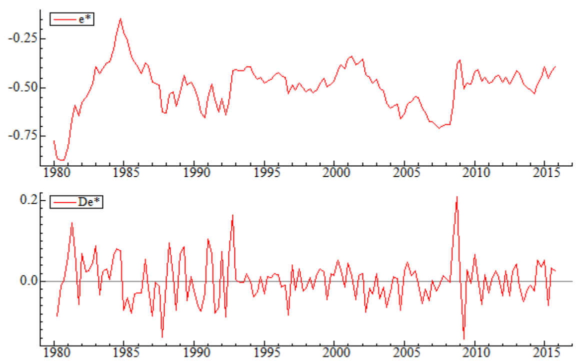

Figure 2.13. Spot GBP/USD Exchange Rate

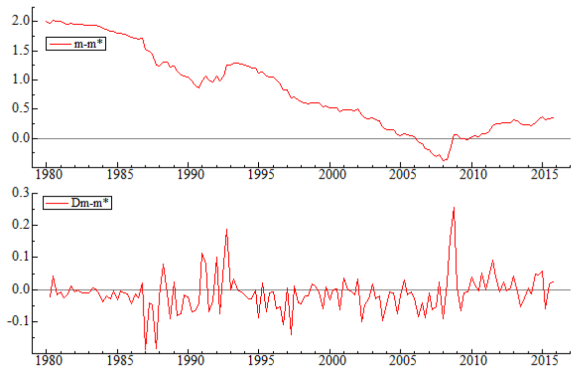

Figure 2.14. Relative M1 for the GBP/USD Model

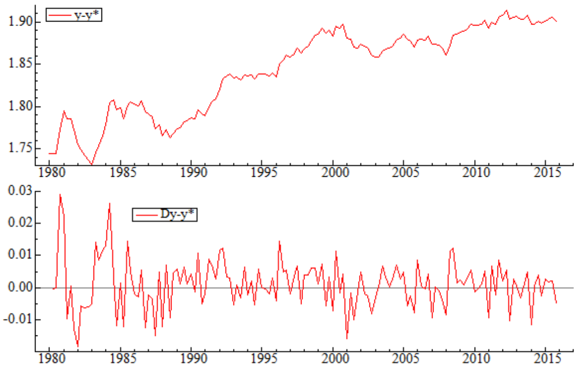

Figure 2.15. Relative GDP for the GBP/USD Model

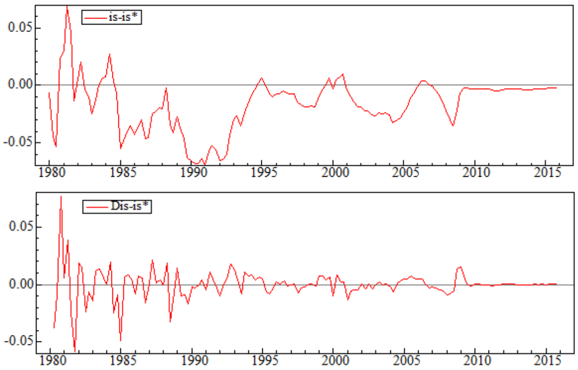

Figure 2.16. Relative Short-Term Interest Rate

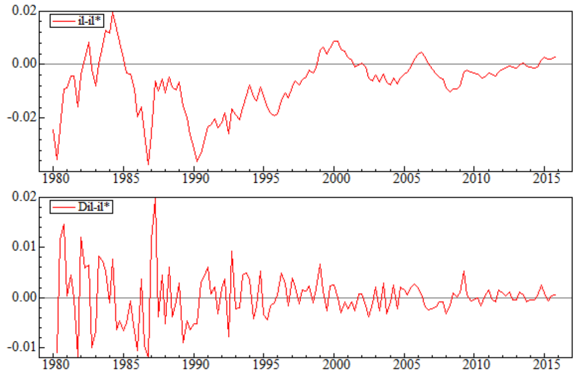

Figure 2.17. Relative Long-Term Interest Rate

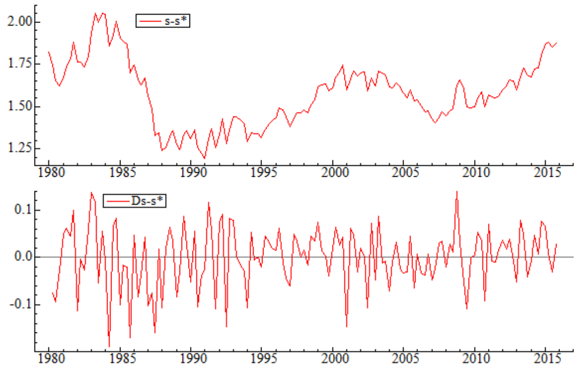

Figure 2.18. Relative Real Stock Prices

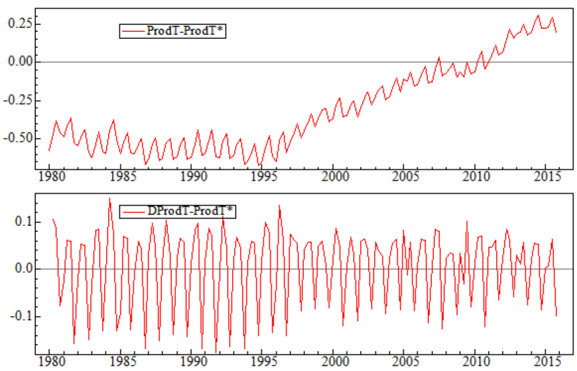

Figure 2.19. Productivity Differential

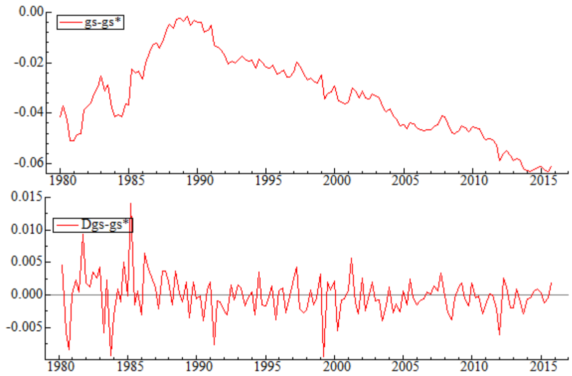

Figure 2.20. Government Spending Differential

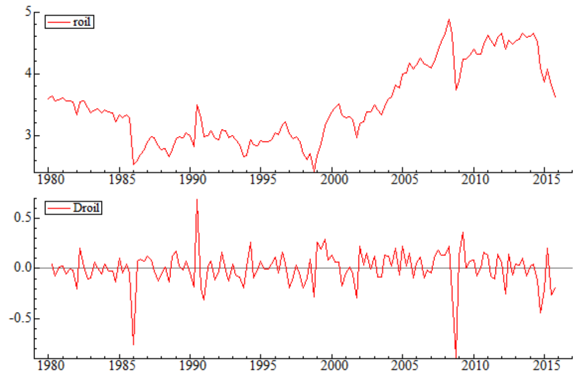

Figure 2.21. The Real Oil Price

Descriptive statistics for the individual variables employed in the RID monetary model and extended model for the USD vis-à-vis the GBP and CAD are presented, respectively, in Tables 2.1 and 2.2., below. It is evident from Tables 2.1. and 2.2. that the variables related to relative income, relative short-term and long-term interest rates, relative productivity and the real oil price for both the GBP/USD and CAD/USD models do not exhibit a normal distribution. Relative share prices and relative government spending for the CAD/USD model, shown in table 2.2., are in addition not normally distributed while the nominal exchange rate and relative money supply for the GBP/USD, shown in Table 2.1., model are also not normally distributed. The finding of non-normality is a result of excess kurtosis, the results of which are also shown in Table 2.1. and Table 2.2., below.

Table 2.1. Descriptive Statistics for the GBP/USD Model Variables

Variable	Mean	Standard Deviation	Skewness	Excess Kurtosis	Normality Test (asymptotically Chi squared)
e*	-0.49	0.11	-0.57	1.30	9.90 [0.00]**
$m - m^*$	0.81	0.68	0.29	-1.1	20.31 [0.00]**
$y - y^*$	1.84	0.05	-0.50	-1.05	38.06 [0.00]**
$i^s - i^{s*}$	-0.01	0.02	-0.18	1.13	9.13 [0.01]*
$i^l - i^{l*}$	-0.00	-0.01	-0.70	0.52	13.16 [0.00]**
$s\text{-}s^*$	1.57	0.19	0.28	-0.39	3.75 [0.15]
$prod^T\text{-}prod^{T*}$	-0.31	0.27	0.59	-0.85	38.76 [0.00]**
$gs\text{-}gs^*$	-0.03	0.01	0.08	-0.80	4.93 [0.08]
roil	3.50	0.62	0.49	-0.93	29.49[0.00]**

Notes: ** and * indicate significance at the 1% and 5% levels, respectively. p-values are in brackets [].

Table 2.2. Descriptive Statistics for the CAD/USD Model Variables

Variable	Mean	Standard Deviation	Skewness	Excess Kurtosis	Normality Test (asymptotically Chi squared)
e*	0.22	0.12	-0.10	-0.74	4.12[0.12]
$m - m^*$	3.38	0.48	-0.29	-0.13	2.58[0.27]
$y - y^*$	2.38	0.04	-0.82	-0.49	59.21[0.00]**
$i^s - i^{s*}$	-0.00	0.01	-0.09	1.21	10.45[0.00]**
$i^l - i^{l*}$	-0.00	0.00	-0.21	-1.02	12.77[0.00]**
$s\text{-}s^*$	2.66	0.27	-0.43	-0.53	11.38[0.00]**
$prod^T\text{-}prod^{T*}$	0.11	0.05	0.28	-0.76	9.02 [0.01]*
$gs\text{-}gs^*$	-0.04	0.00	-0.49	-0.53	15.47[0.00]**
roil	3.50	0.62	0.49	-0.93	29.49[0.00]**

Notes: ** and * indicate significance at the 1% and 5% levels, respectively. p-values are in brackets [].

2.3.2. Tests of Stationarity

A condition for performing cointegration tests is that the series under scrutiny are integrated of order one I(1). By means of an Augmented Dicky-Fuller (ADF) (Dickey and Fuller, 1981) test we show that all the variables in our sample are non-stationary in levels or I(0) and first difference stationary or I(1)). The ADF test results are shown for the GBP/USD and CAD/USD model variables in Tables 2.3. and 2.4., respectively, below. Four lags are specified in the ADF tests as our data our quarterly.

Table 2.3. ADF Unit Root Test Results for the GBP/USD Model Variables

$$\Delta^2 x_t = \pi_0 + \gamma \Delta x_{t-1} + \pi_1 \Delta^2 x_{t-1} + \varepsilon_t$$

Variable	Level		First differences	
	Constant	Constant and trend	Constant	Constant and trend
e*	-2.86	-2.90	-5.81**	-5.79**
$m - m^*$	-1.77	-0.93	-4.40**	-4.66**
$y - y^*$	-0.79	-2.83	-6.20**	-6.18**
$i^s - i^{s*}$	-2.64	-2.94	-7.43**	-7.56**
$i^l - i^{l*}$	-2.74	-2.99	-5.63**	-5.6**
$s\text{-}s^*$	0.09	-1.67	-6.57**	-6.72**
$prod^T\text{-}prod^{T*}$	0.84	-2.79	-4.19**	-4.85**
$gs\text{-}gs^*$	-0.79	-3.33	-4.88**	-5.39**
roil	-1.38	-2.36	-6.45**	-6.44**

Note: the 1% and 5% critical values for the ADF tests are respectively -3.48 and -2.88 and (without trend) and -4.03 and -3.44 (with the trend); setting the lags to be four. ** and * indicate statistical significance at the 1% and 5% levels, respectively.

Table 2.4. ADF Unit Root Test Results for the CAD/USD Model Variables

$$\Delta^2 x_t = \pi_0 + \gamma \Delta x_{t-1} + \pi_1 \Delta^2 x_{t-1} + \varepsilon_t$$

Variable	Level		First differences	
	Constant	Constant and trend	Constant	Constant and trend
e*	-1.86	-1.92	-5.04**	-5.00**
$m - m^*$	-1.51	-1.97	-4.45**	-4.45**
$y - y^*$	-2.34	-1.62	-5.53**	-6.18**
$i^s - i^{s*}$	-2.36	-2.65	-5.55**	-5.53**
$i^l - i^{l*}$	-2.42	-3.43	-5.33**	-5.31**
$s\text{-}s^*$	-1.62	-1.71	-4.13**	-4.09**
$prod^T\text{-}prod^{T*}$	-1.65	-1.70	-4.22**	-4.19**
$gs\text{-}gs^*$	-1.72	-1.52	-5.87**	-5.96**
roil	-1.38	-2.36	-6.45**	-6.44**

Note: the 1% and 5% critical values for the ADF tests are respectively -3.48 and -2.88 (without trend) and -4.03 and -3.44 (with the trend); setting the lags to be four. ** and * indicate statistical significance at the 1% and 5% levels, respectively.

2.3.3. The Econometric Approach

The Johansen (1995) methodology is an offshoot of the Engle and Granger two step methodology for identifying long-run relationship and extends the notion of cointegration developed in Engle and Granger (1987) to develop a coherent methodology to test for and estimate cointegrating relations in a VAR context. Johansen (1988; 1995) introduced the Maximum Likelihood test as a more robust approach to establishing equilibrium relationships amongst variables that are level non-stationary in a multivariate setting. The focus of the methodology is to recognize stochastic trends and cointegrating relationships and examine their composition. Hoover et al. (2008) advocate the approach firstly, because non-stationary economic series are known to be problematic and secondly because they focus on the adaptation of one variable to another and the search for the most favourable outcome for a single factor as well as a congruent structural effect. The Johansen (1995) methodology has been employed in a vast number of studies and is endorsed with regard to monetary models of the exchange rate (see Macdonald, 2007, Chapter 6). Widely recognized are studies conducted by Macdonald and Taylor (1993 and 1994), Sarantis (1994), McNown and Wallace (1994), Chinn (1997a) and Cheung and Chinn (1998). Johansen (1995) assumes an unrestricted vector autoregression model of order p with n I(1) endogenous variables that is driven by a vector X_t with (n x 1) Gaussian errors expressed as an error correction as follows:

$$\Delta X_t = \Pi X_{t-1} + \Gamma_1 \Delta X_{t-1} + ... + \Gamma_p \Delta X_{t-(p-1)} + \gamma D_1 + \varepsilon_1, \qquad (2.25)$$

where X_t is an (n × 1) vector of variables associated with the theoretical long-run relation originating from monetary models, D_1 is a vector of constants, seasonal dummies, and impulse dummies; Γ_i (i =1 ,....., p-1) are (n × n) variable matrices encapsulating the short-run dynamics between model variables, and Π is an (n × n) matrix that is structured as

$\mathbf{\Pi}=\boldsymbol{\alpha}\boldsymbol{\beta}'$, where $\boldsymbol{\alpha}$ denotes the rate of adjustment and $\boldsymbol{\beta}$ matrices denote elements with a long-run trend.

The trace test and maximum eigenvalue statistic are constituents of the maximum likelihood tests introduced by Johansen (1995) to examine cointegration and establish the rank order r of the long-run matrix $\mathbf{\Pi}$. For the purpose of this analysis the trace test is implemented to establish the rank order r of Π. The Johansen (1995) test follows a superlative succession and begins with the null hypothesis of zero cointegrating vectors r=0 against the alternative of at least a single cointegrating vector r ≤1 and progresses with the elimination of supplementary ranks of cointegration until r = i against the alternative r ≤ i+1; the succession halts at r = i as the null can no longer be rejected. Hubrich et al. (2001) find that imposing restrictions by cointegrating rank is ineffectual for forecasts at short intervals however Engle and Yoo (1987) and Clements and Hendry (1998) find that employing an unrestricted VAR at long-horizons yields powerful results. The trace test as expressed in terms of eigenvalues (λ_i) and sample size (T) is represented as follows:

$$\lambda_{trace} = -T\sum_{i=1}^{n}(1-\lambda_i). \tag{2.26}$$

The success of the Johansen (1995) test lies with the sound establishment of the unexpressed VAR as the Johansen (1995) test for cointegration is founded on an unrestricted VAR (Johansen, 1995). Lag length is most frequently determined by information criteria namely the Bayesian information criterion, Schwarz criterion or Akaike information criterion. Burke and Hunter (2005) argue that inadequate choice of lag length can misrepresent the size of the trace test with respect to the null distribution so suggest amplifying the number of lags when serial correlation is detected. We however, do not need to add lags as both exchange rate models are well specified with no serial correlation amongst the residuals (see misspecification tests in Appendix A2.2., Tables A2.4.1. for the GBP/USD model variables and A2.4.2. for the CAD/USD model variables)

Dummy variables are incorporated in both monetary models to counterbalance the effect of exogenous shocks pertaining to the early 1980's global recession, the Big Bang (1986:Q4, 1987;Q1), the 1987 stock market crash (1987:Q4), the Persian Gulf War (1990:Q3), Black Wednesday (1992:Q4), the height of the 2007-2008 subprime crisis and its aftermath (2008:Q4, 2009:Q2, 2009:Q3)[9] and collapse of the oil price (2015:Q1). Other magnified observations are associated with interest rate volatility in the early 1980's.

We employ the procedure described above to distinguish a persistent long-run equilibrium relation in the RID and MRID monetary model for implementation of a specific-to-general approach (Juselius and MacDonald, 2004). Eq. (2.25) is used to estimate the standard RID model as well as the modified RID model, expressed as the vectors in the corresponding rows:

$$X'_{(RID)t} = [e_t^*, m_t - m_t^*, y_t - y_t^*, i_t^s - i_t^{s*}, i_t^l - i_t^{l*}],$$

$$X'_{(MRID)} = [e_t^*, m - m_t^*, y_t - y_t^*, i_t^s - i_t^{s*}, i_t^l - i_t^{l*}, s_t - s_t^*, prod_t^T - prod_t^{T*}, gs_t - gs_t^*, roil_t]$$

The aim of this analysis is to identify elements with a deterministic trend and to substantiate the long-run relation implicit in the model. First root properties of the individual series are investigated, and the series are expected to be non-stationary in levels. The Johansen (1995) procedure is subsequently employed in order to identify a meaningful long-run relationship. Tests of LE (Juselius, 1995) and WE (Johansen, 1992b) are applied to

[9] Impulse dummies were detected by the impulse indicator saturation (IIS) function in Oxmetrics (Hendry and Santos, 2005; Hendry et al., 2008 and Doornik et al., 2013) function in Oxmetrics©. The 2008:Q3 dummy variable accounts for collapse of the world's banking system and the global stock market crash on 29 September 2008 while the 2008:Q4 dummy variable accounts for the period following the height of the subprime mortgage crisis. The 2009:Q2 and 2009:Q3 dummy variables account for the global economic downturn, news of losses in industry and commerce and continued fiscal stimulus by governments worldwide.

ensure soundness of the model as a long-run explanation of the exchange rate. Variables that are found to be weakly exogenous are perceived to propel the system rather than adjust to it in the long-run. Burke and Hunter (2005) claim that recognition of long-run excluded and weakly exogenous variables not only aids in identification of a long run relation but provides insight into the dynamics of such a relation.

2.3.4. The Heterogeneous Case for Unit-Root Testing

To test for stationarity we apply the Augmented Dickey-Fuller test for stationarity to each variable in our sample allowing for a maximum of four lags considering that the data is quarterly. The ADF test assumes the following representation:

$$\Delta\bar{x}_{it} = \mu_i + \psi_i \bar{x}_{it-1} + \sum_{j=1}^{p-1} \psi_j \Delta\bar{x}_{it-j} + \varepsilon_{it} \tag{2.27}$$

where $\bar{x}_{\mathrm{it}} = x_{\mathrm{it}} - \sum_{\mathrm{j=1}}^{\mathrm{t}} \frac{x_{ij}}{t}$

The large sample distribution of the ADF test is not sensitive to the inclusion of differenced time series and controls for the initial values suggesting that each series is homogeneous relative to the point of departure. Eq (2.27) is estimated by OLS and critical values for comparison of Eq. (2.27) are calculated under the null ($\psi_i=0$). The variance is corrected to account for more heterogeneity as well as the model estimated by autoregressive conditional heteroskedasticity (ARCH)[10].

[10] Beirne et al. (2007) show by application of an ADF test where the t-statistics employ White standard errors for all but two countries in their sample, where the ARCH method is employed, that eight out of the ten countries in the sample exhibit stationary real exchange rates at a nominal 5% significance level. The standard errors for Luxemburg and Portugal are extracted from a model that estimates ARCH(1) and ARCH(4) terms in the variance. The ARCH(1) and ARCH(4) estimates are considered more efficient and thus the test statistics more exact on condition that the volatility is consistently well behaved. Considering this the traditional large sample or asymptotic critical values may be adopted.

2.3.5. Tests of LE, WE and Stationarity

The analyses in sections 2.4.1 and 2.4.2. employ asymptotically chi-squared (Johansen, 1992b; Asteriou and Hall, 2015), tests of LE, WE and stationarity for the standard RID model and extended version for the GBP/USD and CAD/USD exchange rates. LE is examined by placing a zero restriction on the pertinent variable of β. Normalisation of the long-run relation on the variable of β is conditional upon rejection of the zero restriction. WE is examined by placing a zero restriction on pertinent variables of α. A variable is considered weakly exogenous if the zero restriction is not rejected thus the variable steers the system rather than adjusting to a long-run equilibrium. In a multivariate framework, the null hypothesis of stationarity is examined by setting each variable in a single cointegrating vector to unity and the remaining variables to zero.

2.4. Empirical Results

2.4.1. Results of the RID Monetary Models

The monetary model depends on stable money demand relations and the assumption that PPP holds. Using series that are I(1), we can observe an exchange rate equation by finding a cointegrating relation and showing via a likelihood ratio test that this variable is neither long-run excluded (Juselius, 1995), nor weakly exogenous (Johansen, 1992b). According to Burke and Hunter (2005), such a finding can help in interpreting and identifying a long-run relation.

Having shown in section 2.3.2 (Tables 2.3. and 2.4.) that the variables in our model are first difference stationary we continue with an empirical test of the RID monetary model of the exchange rate. The VAR model is therefore built upon the $X'_{(RID)t}$ vectors. In order to select the proper lag length, we refer to a preliminary VAR of 4 lags for both exchanges rates and reflect on the significance of F-tests on retained regressors whereby we conclude that the appropriate number of lags for the RID model for the GBP/USD and the CAD/USD is indeed 4. The VAR specification for the standard RID models of the GBP/USD and CAD/USD incorporate a constant, 3 seasonal dummy variables[11] and various impulse dummy variables (described in section 2.3.3.). Misspecification tests of the VAR for the GBP/USD and CAD/USD RID models are reported in Appendix A2.1. (see Tables A2.1. and A2.2., respectively).

The misspecification tests illustrate that at the 5% significance level, both the GBP/USD and the CAD/USD models are free from serial correlation, by means of LM tests of order (8), and ARCH effects, by means of ARCH tests of order (8). The normality tests show that all the variables in our model as well as the vector for the GBP/USD and CAD/USD are normal at a significance level of 5%. We gather that the GBP/USD model is well specified however the CAD/USD model exhibits autocorrelation in the vector.

To this end, we present the estimated eigenvalues and trace statistics of the standard RID monetary model of the GBP/USD and CAD/USD exchange rates in Tables 2.5. and 2.6., respectively, as follows:

Table 2.5. Johansen (1995) Cointegration Test Results for the GBP/USD RID Model
The system comprises of [e^*, $m - m^*$, $y - y^*$, $i^s - i^{s*}$, $i^l - i^{l*}$]

(p –r)	*rank*	Eigenvalue	Trace Test	99% Critical Value	*p-value*
5	$r = 0$	0.26	80.87	76.07	[0.00] **
4	$r \leq 1$	0.13	39.63	54.46	[0.24]
3	$r \leq 2$	0.07	19.39	35.65	[0.48]
2	$r \leq 3$	0.04	8.57	20.04	[0.41]
1	$r \leq 4$	0.02	2.31	6.65	[0.13]

Notes: *r* denotes the number of cointegrating vectors. ** Indicates statistical significance at the 1% level. P-values in brackets [ ].

[11] 3 Seasonal dummy variables as oppose to 4 are employed in our modal to avoid the dummy variable trap. We drop one seasonal dummy variable to prevent multicollinearity. Variables that are highly correlated with each other would not allow for identification of the model parameters.

Table 2.6. Johansen (1995) Cointegration Test Results for the CAD/USD RID Model The system comprises of [e*, $m - m^*$, $y - y^*$, $i^s - i^{s*}$, $i^l - i^{l*}$]					
(p –r)	*rank*	Eigenvalue	Trace Test	99% Critical Value	*p-value*
5	*r = 0*	0.3	96.29	76.07	[0.00] **
4	*r ≤ 1*	0.15	46.74	54.46	[0.06]
3	*r ≤ 2*	0.08	23.18	35.65	[0.25]
2	*r ≤ 3*	0.05	11.72	20.04	[0.17]
1	*r ≤ 4*	0.03	4.73	6.65	[0.03]
Notes: *r* denotes the number of cointegrating vectors. ** Indicates statistical significance at the 1% level. P-values in brackets [ ].					

We observe from Tables 2.5. and 2.6. a single cointegrating vector among the model variables because the null hypothesis of no cointegration is rejected for both RID monetary models (Frankel, 1979a). Since the GBP/USD and CAD/USD money equations appear to contain a single cointegrating relation where rank=1, we estimate the cointegrating vector after normalising on the exchange rate as is indicated in Tables 2.7. for the GBP/USD model and 2.8. for the CAD/USD model. A cointegrating vector is thought to describe a linear combination of nonstationary series and therefore negate normalisation on a stationary variable (Burke and Hunter, 2005, ch5). Furthermore, the exchange rate, as the variable of primary concern is expected to be endogenous. Boswijk (1996) has however advised that supplementary rank conditions be employed to test the soundness of identifying restrictions.

Table 2.7. The Estimated Cointegrating Vectors of the RID Model of the GBP/USD Exchange Rate				
	m-m*	y-y*	$i^s - i^{s*}$	$i^l - i^{l*}$
Coefficient	0.25	2.83	2.33	0.13
Standard error	0.05	0.75	1.23	2.22
t-stat	5.55**	3.78**	1.9*	0.06
Notes: ** and * indicate significance at the 1% and 5% levels, respectively based on standard inference.				

Table 2.8. The Estimated Cointegrating Vectors of the RID Model of the CAD/USD Exchange Rate				
	m-m*	y-y*	$i^s – i^{s*}$	$i^l – i^{l*}$
Coefficient	-0.67	4.80	-18.28	-18.72
Standard error	0.18	1.5	7.13	14.21
t-stat	-3.63**	3.21**	-2.57**	-1.32
Notes: ** and * indicate significance at the 1% and 5% levels, respectively based on standard inference.				

It is important to note that in our monetary model the nominal exchange rate is denoted in terms of the foreign country so that a coefficient rise for the foreign currency would indicate an appreciation of the USD as compared to a depreciation. Upon review of the cointegrating vectors, where the short-interest rate and long-interest rate are taken to be weakly exogenous, we note that the coefficient on the money supply differential is positive (β_1>0) and significant (at the 1% level) for the GBP/USD model while it is negative (β_1<0) for the CAD/USD model (at the 1% level). A coefficient rise in the relative money supply indicates an appreciation of the domestic currency (here the USD) for the GBP/USD model. This result is far from unity and opposes the theory that an increase in domestic money supply relative to the foreign counterpart raises domestic prices and induces a one for one depreciation of the domestic currency. In both models the coefficient on the long-run interest rate differential is insignificant whereas the coefficient on the short-run interest rate differential is positive and significant at the 5% level for the GBP/USD model while it is negative while significant at the 1% level for the CAD/USD model. So, in the case of the GBP/USD model supports the theory that an increase in capital inflows toward the domestic economy relative to a foreign counterpart brings about an appreciation of the domestic currency. The coefficient on the long-run interest rate differential is positive for the GBP/USD model evidence of a wealth effect while the opposite is true for the CAD/USD evidence of a substitution effect. A positive coefficient for the long-run interest rate supports the theory that a decline in the expected rate of domestic inflation heightens the demand for the domestic currency relative to a foreign counterpart and induces an appreciation of the domestic currency. These models do not appear to be models of the nominal exchange rate as several variable coefficients are of the wrong sign and magnitude. In addition, the WE test of the exchange rate in Table 2.10. see Panel B below, indicate that the CAD/USD exchange rate operates like a random walk (Engel and West, 2005).

Tables 2.9. and 2.10. show results of asymptotically chi-squared (Johansen, 1992b; Asteriou and Hall, 2015), tests of LE, WE, and stationarity for the GBP/USD and CAD/USD, respectively. LE is examined by placing a zero restriction on the pertinent variable of β. Normalisation of the long-run relation on the variable of β is conditional upon rejection of the zero restriction. WE is examined by placing a zero restriction on pertinent variables of α. A variable is considered weakly exogenous if the zero restriction is not rejected thus the variable steers the system rather than adjusting to a long-run equilibrium.

The LE tests for the GBP/USD model imply that the short-term and long-term interest rate differentials can be excluded from the cointegration framework thus we infer that the GBP/USD model is misspecified as a long-run model of the exchange rate. For the CAD/USD model the LE tests suggest that the nominal exchange rate, relative money supply and the relative long-run interest rate can be excluded from the model. Therefore, a long-run model founded on the exchange rate may be misspecified, as the corresponding variable has an impact that is statistically not dissimilar from zero.

WE is rejected for the exchange rate and relative income in the GBP/USD model. WE is rejected for the GBP/USD nominal exchange rate so when compared with conventional identification of econometric systems it is considered appropriate to normalise on this variable however WE is not rejected for the CAD/USD nominal exchange rate so in this case one would condition on the exchange rate. Therefore, a model founded on the CAD/USD exchange rate is not possible. The WE test for the CAD/USD model suggests that the exchange rate is primarily a random walk and that it does not adjust to the long run equilibrium but drives the system instead (Hunter, 1992). Engle and West (2005) suggest that exchange rates can help forecast fundamentals. The finding of WE for the relative long-run exchange rate differential in the CAD/USD model is like that of Hunter and Ali (2014) in their USD/JPY RID model. In the case of the CAD/USD model the rejection of WE for the short-term interest rate differential as compared to the finding of WE for the long-term interest rate differential negates the term structure of interest rates since the direction of transmission is from the long-term interest rate to the short-term interest rate and not vice versa. Juselius and Macdonald (2004) find similar evidence of inverse transmission and describe the long-term interest rates as leading the short-term rates in a parity model for the USD/JPY.

The LE tests for the CAD/USD RID model indicate that the nominal exchange rate on which the system has been conditioned can be excluded from the cointegrating framework. Therefore, based on the LE test a long-run model of the exchange rate is not possible. The long-run model would also be incoherent if the impact of the related variables were not distinguishable from zero. The LE test result confirms that it is not feasible to detect using this sample and specification a model of the CAD/USD exchange rate. The conditioning may be reversed in the case of the CAD/USD model if relative share prices, relative government spending, productivity differential and the real oil price are added to the system. The stationarity tests in the multivariate framework corroborate that all the variables in the structure are in levels nonstationary.

Table 2.9. LE, WE, and Stationarity (S) Tests for the GBP/USD RID Model					
	e*	m-m*	y-y*	$i^s - i^{s*}$	I^l-i^{l*}
Panel A: LE tests					
$X^2(1)$	15.28	11.62	9.15	2.71	0.00
p-value	[0.00]**	[0.00]**	[0.00]**	[0.10]	[0.96]
Panel B: WE tests					
$X^2(1)$	12.85	2.54	7.16	0.35	2.30
p-value	[0.00]**	[0.11]	[0.01]**	[0.56]	[0.13]
Panel C: S tests					
$X^2(1)$	32.41	34.69	29.37	20.81	30.3
p-value	[0.00]**	[0.00]**	[0.00]**	[0.00]**	[0.00]**
Notes: ** and * indicate significance at the 1% and 5% levels, respectively. p-values are in brackets [].					

Table 2.10. LE, WE, and Stationarity (S) Tests for the CAD/USD RID Model					
	e*	m-m*	y-y*	$i^s - i^{s*}$	$i^l - i^{l*}$
Panel A: LE tests					
$X^2(1)$	1.85	3.84	4.42	6.45	1.33
p-value	[0.17]	[0.05]	[0.04]*	[0.01]*	[0.25]
Panel B: WE tests					
$X^2(1)$	1.01	0.08	1.09	22.78	0.00
p-value	[0.32]	[0.78]	[0.3]	[0.00]**	[0.99]
Panel C: S tests					
$X^2(1)$	42.11	38.06	20.45	31.93	34.51
p-value	[0.00]**	[0.00]**	[0.00]**	[0.00]**	[0.00]**
Notes: ** and * indicate significance at the 1% and 5% levels, respectively. p-values are in brackets [].					

Considering monetary theory, we conclude that the RID monetary model of the GBP/USD and CAD/USD exchange rate is an incoherent framework for modelling the long-run equilibrium exchange rate since the estimated model variables appear to be of the wrong sign and magnitude.

2.4.2. Results of the Augmented Monetary Models

Our investigation of the standard RID model leads us to query its' validity as a credible model of the long-run exchange rate. We have discussed factors contributing to the failure of the standard monetary model namely, the breakdown of PPP and stable money demand. We examine the implications of cross-country productivity differentials as proposed by Balassa (1964) and Samuelson (1964) and discuss the absence of a measure for real economic activity. Friedmann (1988) introduced the notion of incorporating equity prices that capture long-run economic trends and expectations about the future that were prior to the inclusion of prices immeasurable. Other studies of the monetary models of exchange rate determination have attempted to account for terms of trade and external money shocks by incorporation of a real oil price. We reflect on the factors contributing to past breakdown of the conventional monetary model and continue for this reason to construct and examine a modified monetary model of the exchange rate. The VAR for our modified monetary model rests on the $X'_{(MRID)t}$ vector.

Following the methodology of Hunter and Ali (2014) we assume that the oil price is exogenous to the system since as compared to the other variables it is not relative. We do however, as Hunter and Ali (2014) perform a test of long-run WE to verify this assumption. Johansen and Juselius (1992), Hamilton (1985), Amano and van Norden (1998) and Kim and Roubini (2000) have all in addition conducted studies under the assumption that the oil price is exogenous since shocks to the oil price originate from the Middle East and are therefore exogenous to the macroeconomic architecture under investigation. Tests of LE and WE for the oil price are of particular interest in the CAD/USD model considering that Canada is a major oil producer and that in 2014, 97% of its' oil exports went to the United States.[12] WE tests for the hybrid model of the CAD/USD exchange rate (Table 2.10. Panel B.) however demonstrate that the real oil price is not weakly exogenous with respect to the

[12] See Natural Resources Canada http://www.nrcan.gc.ca/energy/oil-sands/18078.

long-run cointegrating relation, diverging from Amano and van Norden (1998), Kim and Roubini (2000) and Hunter and Ali (2014).

The Choice of lag length was made by reference to the p-values of F-tests on retained regressors. A lag length of 4 is indicated for the VAR specification of the augmented RID model for the GBP/USD as well as for the CAD/USD. The VAR specification for both the augmented RID models incorporate a constant, 3 seasonal dummies and the impulse dummy variables[13] incorporated in the standard RID monetary variable as well as various others that account for exogenous shocks to the relative share price, relative government spending ratio, productivity differential and real oil price (collapse of the oil price in 2015:Q1).

Misspecification tests for the augmented RID monetary model are presented in Appendix A2.1. (see Tables A2.3. for the GBP/USD model and A2.4. for the CAD/USD model). The misspecification tests show that the GBP/USD and CAD/USD models are free from serial correlation at a significance level of 5% and that there is no detectable autoregressive conditional heteroscedasticity amongst the variable residuals. Non-normality is rejected at a significance level of 5% for the individual equations and the system for both exchange rate models. Our misspecification tests indicate that the augmented model is well formed.

The Johansen (1995) cointegration test results for the augmented RID monetary models for the GBP/USD and CAD/USD are presented in Tables 2.11. and 2.12. respectively. The test results show that the null hypothesis of no-cointegration is rejected, where there is confirmation of a single cointegrating vector at the 1% level for both augmented RID models. Evidence of a cointegrating vector in the hybrid models confirms that the added

[13] The impulse dummy variables were detected by the impulse indicator saturation (IIS) function in Oxmetrics (Hendry and Santos, 2005; Hendry et al., 2008 and Doornik et al., 2013) function in Oxmetrics©.

variables adhere to the same stochastic trend as that of the nominal exchange rate and monetary fundamentals incorporated in the standard RID monetary model. Therefore, both augmented RID monetary models are credible equilibrium models.

Table 2.11. Johansen (1995) Cointegration Test Results for the GBP/USD Hybrid Model

The System comprises of [e*, m – m*, y – y*, i^s – i^{s*}, i^l – i^{l*}, s-s*, $prod^T$-$prod^{T*}$, gs-gs*, roil]

(p –r)	*rank*	Eigenvalue	Trace Test	99% Critical Value	p-value
8	*r = 0*	0.44	239.99	204.95	[0.00] **
7	*r ≤ 1*	0.29	158.30	168.36	[0.06]
6	*r ≤ 2*	0.23	110.93	133.57	[0.28]
5	*r ≤ 3*	0.20	74.4	103.18	[0.57]
4	*r ≤ 4*	0.14	42.52	76.07	[0.9]
3	*r ≤ 5*	0.09	21.87	54.46	[0.97]
2	*r ≤ 6*	0.04	9.31	35.65	[0.99]
1	*r ≤ 7*	0.02	3.06	20.04	[0.96]
0	*r≤ 8*	0.00	0.00	6.65	[0.98]

Notes: *r* denotes the number of cointegrating vectors. ** indicates statistical significance at the 1% level. P-values in brackets [ ].

Table 2.12. Johansen (1995) Cointegration Test Results for the CAD/USD Hybrid Model

The System comprises of [e*, m – m*, y – y*, i^s – i^{s*}, i^l – i^{l*}, s-s*, gs-gs*, $prod^T$-$prod^{T*}$, s-s*, roil]

(p –r)	*rank*	Eigenvalue	Trace Test	99% Critical Value	p-value
9	*r = 0*	0.49	254.48	204.95	[0.00]**
8	*r ≤ 1*	0.3	160.88	168.36	[0.04]
7	*r ≤ 2*	0.26	111.14	133.57	[0.27]
6	*r ≤ 3*	0.17	68.65	103.18	[0.77]
5	*r ≤ 4*	0.12	42.44	76.07	[0.9]
4	*r ≤ 5*	0.07	23.91	54.46	[0.94]
3	*r ≤ 6*	0.05	14.370	35.65	[0.82]
2	*r ≤ 2*	0.03	7.23	20.04	[0.56]
1	*r ≤ 7*	0.02	3.32	6.65	[0.07]

Notes: *r* denotes the number of cointegrating vectors. ** indicates statistical significance at the 1% level. P-values in brackets [ ].

The assumption that the added variables do in fact account for features of the real economy as well as productivity differentials is examined by means of LE, WE, and stationarity (S) tests for the hybrid models of the GBP/USD and CAD/USD are presented in Tables 2.13. and 2.14. respectively.

LE tests for the GBP/USD modified RID monetary model, presented in Table 2.13. (see Panel A), show that the nominal exchange rate, relative short-term interest rate government spending ratio and real oil price cannot be excluded from the cointegrating framework at a 5% significance level. LE tests for the CAD/USD hybrid model, presented in Table 2.14 (see Panel A), show that the short and long-run interest rates can be excluded from the cointegrating system at the 5% level of significance. Despite the addition of ancillary variables to the standard RID monetary models the differentials for productivity and share prices can be excluded from the GBP/USD model while in the CAD/USD model LE is rejected for all the variables related to real economic activity and consumption smoothing.

WE tests for the GBP/USD, Table 2.13. (see Panel B), however, demonstrate that WE is rejected for the nominal exchange rate and relative share prices at a 1% significance level. In the case of the GBP/USD model it is in addition evident from Table 2.13. that the differentials related to money supply, income, the long-run interest rate and productivity are exogenous in terms of both weak and strict exogeneity (SE). The inability to reject WE for the differentials relating to productivity, government spending and the real oil price highlights the role of real economic activity and indicates that shocks to the GBP/USD nominal exchange rate do indeed emanate from the economic architecture (Lastrapes, 1992; Chen and Wu, 1997; Enders and Lee, 1997 and An and Kim, 2010).

WE tests for the variables in the CAD/USD hybrid model, Table 2.14. (see Panel B), show that WE is rejected for the nominal exchange rate, relative short-run interest rate and the real oil price. The finding that WE can be rejected for the short-term interest differential but cannot be rejected for the long-term interest rate differential negates the term structure of interest rates since the direction of transmission is generally from the short-term interest rate to the long-term interest rate and not vice a versa. The relative long-run interest rate is in addition shown to be exogenous in terms of both WE and SE. The test result of WE for the oil price opposes that of Amano and van Norden (1998) who argue that shocks to the oil price originate from political unrest in the Middle East so may be viewed as exogenous to the US macroeconomic structure. To differ from Amano and van Norden (1998) we reason that the oil price be considered endogenous to the macroeconomic structure since Canada, as of July 2014, is the world's fifth largest crude oil producer and fourth largest oil exporter with 97% of oil exports going to the United States.[14] Our results support this reasoning as the oil price is shown to be endogenous with respect to the cointegrating relation for the CAD/USD model.

The stationarity tests in the augmented multivariate framework are presented below in Tables 2.13. for the GBP/USD model and 2.14. for the CAD/USD model (see Panel C) and corroborate that the variables in the GBP/USD as well as the CAD/USD structure are level nonstationary, in line with the results of individual unit root tests, presented in section 2.3.2. Table 2.3. and Table 2.4. respectively.

[14] [3]See Natural Resources Canada: Energy Markets Fact Book
http://www.nrcan.gc.ca/sites/www.nrcan.gc.ca/files/energy/files/pdf/2014/14-0173EnergyMarketFacts_e.pdf

Table 2.13. LE, WE, and Stationarity (S) Tests for the GBP/USD Hybrid Model									
	e*	m-m*	y-y*	i^s-i^{s*}	i^l-i^{l*}	s-s*	$prod^T$-$prod^{T*}$	gs-gs*	r_oil
Panel A: LE tests									
$X^2(1)$	11.68	3.75	0.29	21.03	2.50	1.97	1.16	5.57	4.54
p-value	[0.00]**	[0.05]	[0.59]	[0.00]**	[0.11]	[0.16]	[0.28]	[0.02]*	[0.03]*
Panel B: WE tests									
$X^2(1)$	10.64	0.01	1.15	0.75	0.31	20.44	0.06	0.64	0.03
p-value	[0.00]**	[0.91]	[0.28]	[0.39]	[0.58]	[0.00]**	[0.81]	[0.42]	[0.86]
Panel C: S tests									
$X^2(1)$	63.56	58.28	54.23	48.48	65.63	67.41	66.76	67.05	62.74
p-value	[0.00]**	[0.00]**	[0.00]**	[0.00]**	[0.00]**	[0.00]**	[0.00]**	[0.00]**	[0.00]**
Notes: ** and * indicate significance at the 1% and 5% levels, respectively. p-values are in brackets [].									

Table 2.14. LE, WE, and Stationarity (S) Tests for the CAD/USD Hybrid Model									
	e*	m-m*	y-y*	i^s-i^{s*}	i^l-i^{l*}	s-s*	$prod^T$-$prod^{T*}$	gs-gs*	r_oil
Panel A: LE tests									
$X^2(1)$	31.30	20.72	32.91	5.81	5.12	29.72	39.22	18.51	27.70
p-value	[0.00]**	[0.00]**	[0.00]**	[0.02]*	0.02]*	[0.00]**	[0.00]**	[0.00]**	[0.00]**
Panel B: WE tests									
$X^2(1)$	10.14	3.45	0.78	22.9	0.33	0.02	1.98	6.47	11.58
p-value	[0.00]**	[0.06]	[0.38]	[0.00]**	[0.57]	[0.89]	[0.16]	[0.01]*	[0.00]**
Panel C: S tests									
$X^2(1)$	83.60	75.51	88.7	60.3	59.64	59.93	57.3	56.21	59.03
p-value	[0.00]**	[0.00]**	[0.00]**	[0.00]**	[0.00]**	[0.00]**	[0.00]**	[0.00]**	[0.00]**
Notes: ** and * indicate significance at the 1% and 5% levels, respectively. p-values are in brackets [].									

In summation, the incorporation of variables affecting stable money demand and the validity of PPP to the standard RID monetary model of the exchange rate reverses the long run projection for the nominal exchange rate in the case of the CAD/USD model. Reversal of the WE result for the nominal CAD/USD exchange rate attests to adjustment of the long-run evaluation and improves econometric performance of the model, which is a significant advance (Juselius and Macdonald, 2004).

Table 2.15. The Estimated Cointegrating Vector of the Hybrid Model of the GBP/USD Exchange Rate								
	m-m*	y-y*	i^s-i^{s*}	i^l-i^{l*}	s-s*	$prod^T$-$prod^{T*}$	gs-gs*	r_oil
Coefficient	-0.15	0.48	-6.72	3.42	-0.37	-0.48	-4.81	0.19
Standard errors	0.05	0.75	1.23	2.27	0.15	0.19	3.13	0.06
t-stat	-3.15**	0.64	-5.45**	1.51	-2.48*	-2.47*	-1.54	3.00**
Notes: ** and * indicate significance at the 1% and 5% levels, respectively based on standard inference.								

Table 2.16. The Estimated Cointegrating Vectors of the Hybrid Model of the CAD/USD Exchange Rate								
	m-m*	y-y*	i^s-i^{s*}	i^l-i^{l*}	s-s*	$prod^T$-$prod^{T*}$	gs-gs*	r_oil
Coefficient	-0.33	-3.00	1.80	-3.17	-0.49	1.69	-7.81	-0.16
Standard errors	0.06	0.33	0.67	1.27	0.04	0.15	1.2	0.02
t-stat	-5.88**	9.17**	2.7**	-2.5*	-13.32**	11.27**	-6.56*	-6.79**
Notes: ** and * indicate significance at the 1% and 5% levels, respectively based on standard inference.								

The estimated cointegrating vectors, normalised on the exchange rate, are presented for the GBP/USD and CAD/USD modified RID models in Tables 2.15. and 2.16., respectively. Boswijk (1996) states that supplementary rank conditions are a prerequisite for empirical identification. On the other hand, Burke and Hunter (2005, ch 5) assert that disqualifying normalisation on variables that are either long-run excluded or weakly exogenous is a sound approach for empirical identification. Normalisation is not applicable to stationary variables since cointegration is a characteristic of two or more non-stationary variables.

The estimated vectors for the GBP/USD indicate that the coefficients on the real oil price is positive (Johansen and Juselius, 1992) in both models so trigger an appreciation of the domestic currency, US dollar. The variable coefficients in both models, apart from relative income, the long-run interest rate differential and relative government spending in the GBP/USD model, are significant at the 1% level. Several coefficients are however of the wrong sign and magnitude. The money supply differential in both models is of the correct sign but are different from unity so a one-for-one depreciation of the domestic currency does not occur. The coefficient on the long-run interest rate in the CAD/USD model is slightly larger than that of the short-term interest rate differential, a finding consistent with Frankel's (1979a) theory. The coefficient on the share price in both models is negative so that the substitution effect governs money demand functions between the UK, Canada and the USA (Friedman, 1988; McCornac, 1991; Caruso, 2001) while the differentials for government spending have a sign opposing theory so prompt a depreciation of the domestic currency. In light of the estimated vector and coefficients that are either insignificant or of the wrong sign or magnitude the modified RID model is validated in the next section.

2.4.3. Validation of the Modified Monetary Model of the Exchange Rate

Our analysis in section 2.4.1 articulates the deficiency of the standard RID model (Frankel, 1979a) when applied to the GBP/USD and CAD/USD exchange rates. We ascribe this deficiency to the breakdown of purchasing power parity and stable money demand functions. To enhance the RID model, we incorporate variables that capture real economic activity as well as long-run dynamics and formulate the modified model. We follow Hunter and Ali (2014) and pursue a general-to-specific approach (Davidson et al, 1978 and Hendry and Mizon, 1993) dependent upon the outcome of LE and WE tests. This approach is adopted to construct a powerful yet compact specification of the long-term relationship between the nominal exchange rate and real economic fundamentals. We identified a single cointegrating vector of rank r=1 in both exchange rate models and recognise the following framework pertaining to the α and β vectors:

$$e \quad \nabla m \quad \nabla y \quad \nabla i^s \quad \nabla i^l \quad \nabla prod^T \quad \nabla gs \quad roil$$

$$\beta' = [-1 \quad \beta_2 \quad \beta_3 \quad \beta_4 \quad \beta_5 \quad \beta_6 \quad \beta_7 \quad \beta_8 \quad \beta_9],$$

$$\beta' = [-1 \quad 0 \quad \beta_3 \quad \beta_4 \quad \beta_5 \quad \beta_6 \quad \beta_7 \quad \beta_8 \quad \beta_9],$$

$$\beta' = [-1 \quad \beta_2 \quad 0 \quad \beta_4 \quad \beta_5 \quad \beta_6 \quad \beta_7 \quad \beta_8 \quad \beta_9],$$

$$\beta' = [-1 \quad \beta_2 \quad \beta_3 \quad 0 \quad \beta_5 \quad \beta_6 \quad \beta_7 \quad \beta_8 \quad \beta_9],$$

$$\beta' = [-1 \quad \beta_2 \quad \beta_3 \quad \beta_4 \quad 0 \quad \beta_6 \quad \beta_7 \quad \beta_8 \quad \beta_9],$$

$$\beta' = [-1 \quad \beta_2 \quad \beta_3 \quad \beta_4 \quad \beta_5 \quad 0 \quad \beta_7 \quad \beta_8 \quad \beta_9],$$

$$\beta' = [-1 \quad \beta_2 \quad \beta_3 \quad \beta_4 \quad \beta_5 \quad \beta_6 \quad 0 \quad \beta_8 \quad \beta_9],$$

$$\beta' = [-1 \quad \beta_2 \quad \beta_3 \quad \beta_4 \quad \beta_5 \quad \beta_6 \quad \beta_7 \quad 0 \quad \beta_9] \text{ and}$$

$$\beta' = [-1 \quad \beta_2 \quad \beta_3 \quad \beta_4 \quad \beta_5 \quad \beta_6 \quad \beta_7 \quad \beta_8 \quad 0]$$

Where ∇ represents the differential between the domestic (US) and foreign variables (UK and Canada). To normalise on the exchange rate, we impose β_1=-1. From tests of LE and WE for the GBP/USD model we impose restrictions of strict exogeneity (SE) for the relative money supply (β_2=0 and α_2=0), the relative income (β_3 =0 and α_3=0), the relative long-run interest rate (β_5=0 and α_5=0) and for relative productivity (β_7=0 and α_7=0). We also impose β_6=0 for the relative share price as this variable can be long-run excluded but is not weakly exogenous. From the LE tests for the CAD/USD we impose restrictions of SE for the relative long-run interest rate (β_5 =0 and α_5=0). We also impose β_4 =0 for the relative short-run interest rate that can be long-run excluded but is not shown to be weakly exogenous.

Further:

$$e \quad \nabla m \quad \nabla y \quad \nabla i^{s} \quad \nabla i^{l} \quad \nabla prod^{T} \quad \nabla gs \quad roil$$

$$\alpha' = [\alpha_1 \quad \alpha_2 \quad \alpha_3 \quad \alpha_4 \quad \alpha_5 \quad \alpha_6 \quad \alpha_7 \quad \alpha_8 \quad \alpha_9],$$

$$\alpha' = [0 \quad \alpha_2 \quad \alpha_3 \quad \alpha_4 \quad \alpha_5 \quad \alpha_6 \quad \alpha_7 \quad \alpha_8 \quad \alpha_9],$$

$$\alpha' = [\alpha_1 \quad 0 \quad \alpha_3 \quad \alpha_4 \quad \alpha_5 \quad \alpha_6 \quad \alpha_7 \quad \alpha_8 \quad \alpha_9],$$

$$\alpha' = [\alpha_1 \quad \alpha_2 \quad 0 \quad \alpha_4 \quad \alpha_5 \quad \alpha_6 \quad \alpha_7 \quad \alpha_8 \quad \alpha_9],$$

$$\alpha' = [\alpha_1 \quad \alpha_2 \quad \alpha_3 \quad 0 \quad \alpha_5 \quad \alpha_6 \quad \alpha_7 \quad \alpha_8 \quad \alpha_9],$$

$$\alpha' = [\alpha_1 \quad \alpha_2 \quad \alpha_3 \quad \alpha_4 \quad 0 \quad \alpha_6 \quad \alpha_7 \quad \alpha_8 \quad \alpha_9],$$

$$\alpha' = [\alpha_1 \quad \alpha_2 \quad \alpha_3 \quad \alpha_4 \quad \alpha_5 \quad 0 \quad \alpha_7 \quad \alpha_8 \quad \alpha_9],$$

$$\alpha' = [\alpha_1 \quad \alpha_2 \quad \alpha_3 \quad \alpha_4 \quad \alpha_5 \quad \alpha_6 \quad 0 \quad \alpha_8 \quad \alpha_9],$$

$$\alpha' = [\alpha_1 \quad \alpha_2 \quad \alpha_3 \quad \alpha_4 \quad \alpha_5 \quad \alpha_6 \quad \alpha_7 \quad 0 \quad \alpha_9],$$

$$\alpha' = [\alpha_1 \quad \alpha_2 \quad \alpha_3 \quad \alpha_4 \quad \alpha_5 \quad \alpha_6 \quad \alpha_7 \quad 0 \quad \alpha_9] \text{and}$$

$$\alpha' = [\alpha_1 \quad \alpha_2 \quad \alpha_3 \quad \alpha_4 \quad \alpha_5 \quad \alpha_6 \quad \alpha_7 \quad \alpha_8 \quad 0]$$

From the WE tests for the GBP/USD model we impose α_2=0 for the money supply differential, α_3=0 for the income differential, α_4=0 for the short-run interest rate differential, α_5=0 for the long-run interest rate differential, α_7=0 for the productivity differential, α_8=0 for the government spending differential and α_9=0 for the real oil price while from the WE tests for the CAD/USD model we impose α_2=0 for the money supply differential, α_3=0 for the income differential, α_4=0 for the short-run interest rate differential, α_5=0 for the long-run interest rate differential, α_6=0 for the share price differential and α_7=0 for the productivity differential.

Based on this framework we control for long-run exclusions for the GBP/USD and CAD/USD hybrid models by sequentially imposing zero restrictions on the loading factors, α, that are weakly exogenous to the cointegrating relations (see Panel B in Tables 2.13. and 2.14.). The restrictions on weakly exogenous variables are empirically feasible considering the small magnitude of the variable coefficients and established monetary doctrine. Regarding the augmented variables, we refrain from imposing WE on the share price differential for the hybrid GBP/USD model as from the individual tests of WE it is shown to be endogenous to the system. We in addition do not impose a zero restriction on the short-run interest rate differential, government spending differential and real oil price for the GBP/USD model as we deduce from the individual LE tests that these variables are long-run excluded at a significance level of 5%.

Results of the joint tests for the GBP/USD model show that the restrictions imposed on the variables in the GBP/USD hybrid model are firmly accepted for $\beta_2 = 0$, $\beta_3 = 0$ and $\beta_5 = 0$ (see Tables 2.17.-2.19.) but is rejected for $\beta_7 = 0$ at a significance level of 95% (see Tables 2.20. and 2.21.). The restricted long-term relationship for $\beta_3 = 0$ in Table 2.20. and for $\beta_5 = 0$ in Table 2.21. normalised on the exchange rate show that all variables aside from the long-run interest rate differential are significant at the 5% level while the relative money supply and real oil price have their hypothesised signs and a substitution effect is evident from the negative relative share price coefficient. The restricted long-term relationship for $\beta_2 = 0$ in Table 2.19. normalised on the exchange rate show that all variables aside from the share price differential and government spending differential are significant, at the 5% level, in addition all the variables apart from the short-run interest rate differential and productivity differential have their hypothesised signs. These results confirm stability of the hybrid model specification and the importance of real economic variables in determining the behaviour of the nominal exchange rate.

Table 2.17. Exclusion of y-y* for the GBP/USD Hybrid Model

Joint tests of WE and LE conditional on r = 1 in the HM model.

Tests under the null:	Statistics [p-value]
(1) β_3=0	X^2(1) = 0.29 [0.59]
(2) β_3=0, α_3=0	X^2(2) = 1.19 [0.55]
(3) β_3=0, α_3=0, α_2=0	X^2(3) = 1.21 [0.75]
(4) β_3=0, α_3=0, α_2=0, α_9=0	X^2(4) = 1.21 [0.88]
(5) β_3=0, α_3=0, α_2=0, α_9=0 α_7=0	X^2(5) = 2.07 [0.84]
(6) β_3=0, α_3=0, α_2=0, α_9=0, α_7=0, α_5=0	X^2(6) = 2.08 [0.91]
(7) β_3=0, α_3=0, α_2=0, α_9=0, α_7=0, α_5=0, α_8=0	X^2(7) = 2.98 [0.89]
(8) β_3=0, α_3=0, α_2=0, α_9=0, α_7=0, α_5=0, α_8=0, α_4=0	X^2(8) = 4.14 [0.84]

The implied long-run relation by test (8):

	(m-m*)	(i^s-i^{s*})	(i^l-i^{l*})	(s-s*)	($prod^T$-$prod^{T*}$)	(gs-gs*)	(r_oil)
Coef.	-0.18	-6.11	2.66	-0.35	-0.52	-5.80	0.18
Standard error	0.03	1.05	2.08	0.13	0.19	3.06	0.06
t-stat	-6.00**	-5.08**	1.29	-2.69**	-2.74**	-1.89*	3.18**

Notes: ** and * indicate significance at the 1% and 5% levels, respectively based on standard inference. p-values are in square brackets [.].

Table 2.18. Exclusion of i^l-i^{l*} for the GBP/USD Hybrid Model

Joint tests of WE and LE conditional on r = 1 in the HM model.

Tests under the null:	Statistics [p-value]
(1) β_5=0	X^2(1) = 2.50 [0.11]
(2) β_5=0, α_5=0	X^2(2) = 2.60 [0.27]
(3) β_5=0, α_5=0, α_2=0, α_9=0	X^2(3) = 2.71 [0.44]
(4) β_5=0, α_5=0, α_2=0, α_9=0, α_7=0	X^2(4) = 3.06 [0.55]
(5) β_5=0, α_5=0, α_2=0, α_9=0, α_7=0, α_8=0	X^2(5) = 3.7 [0.59]
(6) β_5=0, α_5=0, α_2=0, α_9=0, α_7=0, α_8=0, α_4=0	X^2(6) = 3.91 [0.69]
(7) β_5=0, α_5=0, α_2=0, α_9=0, α_7=0, α_8=0, α_4=0, α_3=0	X^2(7) = 4.45 [0.73]
(8) β_5=0, α_5=0, α_2=0, α_9=0, α_7=0, α_8=0, α_4=0, α_3=0	X^2(8) = 6.22 [0.62]

The implied long-run relation by test (8):

	(m-m*)	(y-y*)	(i^l - i^{l*})	(s-s*)	($prod^T$-$prod^{T*}$)	(gs-gs*)	(r_oil)
Coef.	-0.18	-0.21	-5.28	-0.30	-0.47	-6.22	0.15
Standard error	0.05	0.67	1.02	0.13	0.19	3.00	0.06
t-stat	-4.5**	-0.31	-5.21**	-2.27*	-2.55*	-2.08*	2.63**

Notes: ** and * indicate significance at the 1% and 5% levels, respectively based on standard inference. p-values are in square brackets [.].

Table 2.19. Exclusion of m-m* for the GBP/USD Hybrid Model

Joint tests of WE and LE conditional on r = 1 in the HM model.

Tests under the null:	Statistics [p-value]
(1) β_2=0	X^2(1) = 3.75 [0.05]
(2) β_2=0, α_2=0	X^2(2) = 3.95 [0.14]
(3) β_2=0, α_2=0, α_9=0	X^2(3) = 5.83 [0.12]
(4) β_2=0, α_2=0, α_9=0, α_7=0	X^2(4) = 6.58 [0.16]
(5) β_2=0, α_2=0, α_9=0, α_7=0, α_5=0	X^2(5) = 7.63 [0.18]
(6) β_2=0, α_2=0, α_9=0, α_7=0, α_5=0, α_8=0	X^2(6) = 7.74 [0.26]
(7) β_2=0, α_2=0, α_9=0, α_7=0, α_5=0, α_8=0, α_4=0	X^2(7) = 8.95 [0.26]
(8) β_2=0, α_2=0, α_9=0, α_7=0, α_5=0, α_8=0, α_4=0, α_3=0	X^2(8) = 11.55 [0.17]

The implied long-run relation by test (8):

	(y-y*)	(i^s-i^{s*})	(i^l - i^{l*})	(s-s*)	($prod^T$-$prod^{T*}$)	(gs-gs*)	(r_oil)
Coef.	3.8	-12.50	8.2	-0.23	-1.07	1.23	0.58
Standard error	0.92	2.31	4.37	0.28	0.37	6.00	0.12
t-stat	4.26**	-5.41**	1.9*	-0.78	-3.67**	0.20	5.27**

Notes: ** and * indicate significance at the 1% and 5% levels, respectively based on standard inference. p-values are in square brackets [.].

Table 2.20. Exclusion of $prod^T$-$prod^{T*}$ for the UK Hybrid Model

Joint tests of WE and LE conditional on r = 1 in the HM model.

Tests under the null:	Statistics [p-value]
(1) β_7=0	X^2(1) = 5.57 [0.02]*
(2) β_7=0, α_7=0	X^2(2) = 6.32 [0.04]*
(3) β_7=0, α_7=0, α_3=0	X^2(3) = 8.76 [0.03]*
(4) β_7=0, α_7=0, α_3=0, α_2=0	X^2(4) = 8.78 [0.07]
(5) β_7=0, α_7=0, α_3=0, α_2=0 α_9=0	X^2(5) = 8.96 [0.11]
(6) β_7=0, α_7=0, α_3=0, α_2=0 α_9=0, α_5=0	X^2(6) = 9.33 [0.16]
(7) β_7=0, α_7=0, α_3=0, α_2=0 α_9=0, α_5=0, α_7=0	X^2(7) = 9.90 [0.2]
(8) β_7=0, α_7=0, α_3=0, α_2=0 α_9=0, α_5=0, α_7=0, α_4=0	X^2(8) =11.79 [0.16]

The implied long-run relation by test (8):

	(m-m*)	(y-y*)	(i^s-i^{s*})	(i^l-i^{l*})	(s-s*)	(gs-gs*)	(r_oil)
Coef.	-0.14	-0.1	-4.51	1.52	-0.53	-4.5	0.05
Standard error	0.04	0.62	0.95	1.89	0.12	2.44	0.04
t-stat	-3.5**	-0.16**	-4.79**	0.79	-5.52**	-1.84*	1.15**

Notes: ** and * indicate significance at the 1% and 5% levels, respectively based on standard inference. p-values are in square brackets[.].

Table 2.21. Exclusion of $prod^{T}$-$prod^{T*}$ for the UK Hybrid Model							
Joint tests of WE and LE conditional on r = 1 in the HM model.							
Tests under the null:						Statistics [p-value]	
(1) β_7=0						$X^2(1) = 5.57$ [0.02]*	
(2) β_7=0, α_7=0						$X^2(2) = 6.32$ [0.04]*	
(3) β_7=0, α_7=0, α_3=0						$X^2(3) = 8.76$ [0.03]*	
The implied long-run relation by test (3):							
	(m-m*)	(y-y*)	(i^s-i^{s*})	(i^l-i^{l*})	(s-s*)	(gs-gs*)	(r_oil)
Coef.	-0.15	-0.1	-4.76	2.46	-0.56	-4.93	0.04
Standard error	0.04	0.60	0.91	1.82	0.11	2.35	0.04
t-stat	-3.95**	-0.16	-5.23**	13.61**	-5.09**	-2.09*	2.95**
Notes: ** and * indicate significance at the 1% and 5% levels, respectively based on standard inference. p-values are in square brackets[.].							

Individual tests of LE for the CAD/USD hybrid model show that the relative short and long run interest rates cannot be excluded at the 5% level of significance (Table 2.14., Panel A) however for robustness we run joint tests of LE and WE. For the CAD/USD hybrid model the restriction imposed for β_4 = 0 is firmly rejected at the 95% level of significance (see Table 2.23., below) and the restriction imposed for β_5= 0 is firmly rejected at the 95 % level of significance (see Table 2.24., below). The long-term relationship normalised on the CAD/USD nominal exchange rate depicted in Table 2.16. show that all variables aside from the share price differential and the coefficients on the relative money supply, the relative short-run interest rate and the productivity differential are of the correct sign while the substitution effect is evident from the negative share price differential. These results in addition confirm stability of the hybrid model specification and the importance of real economic variables in determining the behaviour of the CAD/USD nominal exchange rate.

Table 2.22. Exclusion of i^s-i^{s*} for the CAD/USD Hybrid Model							
Joint tests of WE and LE conditional on r = 1 in the HM model.							
Tests under the null:						Statistics [p-value]	
(1) β_4=0						X^2(1) = 5.81 [0.02]*	
(2) β_4=0, α_6=0						X^2(2) = 6.22 [0.05]*	
(3) β_4=0, α_6=0, α_5=0						X^2(3) = 7.84 [0.05]*	
(4) β_4=0, α_6=0, α_5=0, α_3=0						X^2(4) = 8.46 [0.08]	
(5) β_4=0, α_6=0, α_5=0, α_3=0, α_7=0						X^2(5) = 11.25 [0.05]*	
(6) β_4=0, α_6=0, α_5=0, α_3=0, α_7=0, α_2=0						X^2(6) = 21.98 [0.00]**	
The implied long-run relation by test (6):							
	(m-m*)	(y-y*)	(i^l-i^{l*})	(s-s*)	($prod^T$-$prod^{T*}$)	(gs-gs*)	(r_oil)
coefficient	-0.4	-3.25	-3.25	-0.49	1.92	-8.57	-0.18
standard error	0.06	0.35	1.24	0.05	0.19	1.43	0.03
t-stat	-6.35**	-9.29**	-2.63**	-10.54**	9.94**	-6.00**	-6.46**
Notes: ** and * indicate significance at the 1% and 5% levels, respectively based on standard inference. p-values are in square brackets[.].							

Table 2.23. Exclusion of i^s-i^{s*} for the CAD/USD Hybrid Model							
Joint tests of WE and LE conditional on r = 1 in the HM model.							
Tests under the null:						Statistics [p-value]	
(1) B_4=0						X^2(1) = 5.81 [0.016]*	
(2) β_4=0, α_6=0						X^2(2) = 6.22 [0.05]*	
(3) β_4=0, α_6=0, α_5=0						X^2(3) = 7.84 [0.05]*	
The implied long-run relation by test (3):							
	(m-m*)	(y-y*)	(i^l-i^{l*})	(s-s*)	($prod^T$-$prod^{T*}$)	(gs-gs*)	(r_oil)
coefficient	-0.24	-2.42	-1.16	-0.52	1.6	-5.84	-0.12
standard error	0.05	0.27	0.95	0.04	0.15	1.1	0.02
t-stat	-4.6**	-9.30**	-1.22	-14.30**	11.35**	-5.34**	-5.72**
Notes: ** and * indicate significance at the 1% and 5% levels, respectively based on standard inference. p-values are in square brackets[.].							

Table 2.24. Exclusion of i^s-i^{s*} for the CAD/USD Hybrid Model							
Joint tests of WE and LE conditional on r = 1 in the HM model.							
Tests under the null:						Statistics [p-value]	
(1) β_4=0						X^2(1) = 5.81 [0.016]*	
(2) β_4=0, α_6=0						X^2(2) = 6.22 [0.05]*	
(3) β_4=0, α_6=0, α_5=0						X^2(3) = 7.84 [0.05]*	
(4) β_4=0, α_6=0, α_5=0, α_7=0						X^2(5) = 11.04 [0.03]*	
(5) β_4=0, α_6=0, α_5=0, α_7=0, α_2=0						X^2(5) = 21.68 [0.00]**	
The implied long-run relation by test (5):							
	(m-m*)	(y-y*)	(i^l-i^{l*})	(s-s*)	($prod^T$-$prod^{T*}$)	(gs-gs*)	(r_oil)
coefficient	-0.39	-3.2	-2.81	-0.49	1.88	-8.44	-0.18
standard error	0.06	0.34	1.21	0.05	0.19	1.40	0.03
t-stat	-6.22**	-9.38**	-2.31*	-10.75**	9.94**	-6.07**	-6.52**
Notes: ** and * indicate significance at the 1% and 5% levels, respectively based on standard inference. p-values are in square brackets[.].							

Table 2.25. Exclusion of il-il* for the CAD/USD Hybrid Model	
Joint tests of WE and LE conditional on r = 1 in the HM model.	
Tests under the null:	Statistics [p-value]
(1) $\beta_5 = 0$	$X^2(1) = 5.12$ [0.02]*
(2) $\beta_5 = 0, \alpha_5 = 0$	$X^2(2) = 6.35$ [0.04]*
(3) $\beta_5 = 0, \alpha_5 = 0, \alpha_6 = 0$	$X^2(3) = 6.51$ [0.09]
(4) $\beta_5 = 0, \alpha_5 = 0, \alpha_6 = 0, \alpha_3 = 0$	$X^2(4) = 9.01$ [0.07]
(5) $\beta_5 = 0, \alpha_5 = 0, \alpha_6 = 0, \alpha_3 = 0, \alpha_7 = 0$	$X^2(5) = 10.51$ [0.06]
(6) $\beta_5 = 0, \alpha_5 = 0, \alpha_6 = 0, \alpha_3 = 0, \alpha_7 = 0, \alpha_2 = 0$	$X^2(6) = 27.55$ [0.00]**
The implied long-run relation by test (6):	

	(m-m*)	(y-y*)	(i^s-i^{s*})	(s-s*)	($prod^T$-$prod^{T*}$)	(gs-gs*)	(r_oil)
coefficient	-0.41	-3.46	0.78	-0.49	1.84	-9.21	-0.22
standard error	0.07	0.42	0.72	0.05	0.20	1.68	0.03
t-stat	-12.21**	-8.17**	1.08	-9.33**	9.13**	-5.48**	-11.92**

Notes: ** and * indicate significance at the 1% and 5% levels, respectively based on standard inference. p-values are in square brackets [.].

Table 2.26. Exclusion of i^l-i^{l*} for the CAD/USD Hybrid Model	
Joint tests of WE and LE conditional on r = 1 in the HM model.	
Tests under the null:	Statistics [p-value]
(1) $\beta_5 = 0$	$X^2(1) = 5.12$ [0.024]*
(2) $\beta_5 = 0, \alpha_5 = 0$	$X^2(2) = 6.35$ [0.04]*
(3) $\beta_5 = 0, \alpha_5 = 0, \alpha_6 = 0$	$X^2(3) = 6.51$ [0.09]
(4) $\beta_5 = 0, \alpha_5 = 0, \alpha_6 = 0, \alpha_7 = 0$	$X^2(5) = 8.74$ [0.07]
(5) $\beta_5 = 0, \alpha_5 = 0, \alpha_6 = 0, \alpha_7 = 0, \alpha_2 = 0$	$X^2(5) = 22.48$ [0.00]**
The implied long-run relation by test (5):	

	(m-m*)	(y-y*)	(i^s-i^{s*})	(s-s*)	($prod^T$-$prod^{T*}$)	(gs-gs*)	(r_oil)
coefficient	-0.4	-3.37	1.36	-0.48	1.7	-9.13	-0.21
standard error	0.07	0.4	0.68	0.05	0.19	1.57	0.03
t-stat	-6.5**	-8.51**	2.03*	-9.65**	9.05**	-5.81**	6.63**

Notes: ** and * indicate significance at the 1% and 5% levels, respectively based on standard inference. p-values are in square brackets[.].

Table 2.27. Exclusion of i^l-i^{l*} for the CAD/USD Hybrid Model	
Joint tests of WE and LE conditional on r = 1 in the HM model.	
Tests under the null:	Statistics [p-value]
(1) $\beta_5 = 0$	$X^2(1) = 5.11$ [0.02]*
(2) $\beta_5 = 0, \alpha_5 = 0$	$X^2(2) = 6.35$ [0.04]*
(3) $\beta_5 = 0, \alpha_5 = 0, \alpha_6 = 0$	$X^2(3) = 6.51$ [0.1]
(4) $\beta_5 = 0, \alpha_5 = 0, \alpha_6 = 0, \alpha_7 = 0$	$X^2(4) = 8.74$ [0.07]
The implied long-run relation by test (4):	

	(m-m*)	(y-y*)	(i^s-i^{s*})	(s-s*)	($prod^T$-$prod^{T*}$)	(gs-gs*)	(r_oil)
coefficient	-0.24	-2.48	1.04	-0.52	1.47	-5.98	-0.14
standard error	0.05	0.31	0.54	0.04	0.15	1.25	0.03
t-stat	-4.46**	-7.89**	1.94*	-13.32**	23.2**	-4.8**	-5.64**

Notes: ** and * indicate significance at the 1% and 5% levels, respectively based on standard inference. p-values are in square brackets[.].

Table 2.28. Exclusion of i^l-i^{l*} for the CAD/USD Hybrid Model							
Joint tests of WE and LE conditional on r = 1 in the HM model.							
Tests under the null:						Statistics [p-value]	
(1) $\beta_5 = 0$						$X^2(1) = 5.12$ [0.03]*	
(2) $\beta_5 = 0$, $\alpha_5=0$						$X^2(2) = 6.35$ [0.04]*	
The implied long-run relation by test (2):							
	(m-m*)	(y-y*)	(i^s-i^{s*})	(s-s*)	($prod^T$-$prod^{T*}$)	(gs-gs*)	(r_oil)
coefficient	-0.24	-2.53	0.99	-0.51	1.51	-6.32	-0.14
standard error	0.05	0.31	0.53	0.04	0.15	1.21	0.02
t-stat	-4.68**	-8.26**	1.90*	-13.83**	10.34**	-11.09**	-5.88**
Notes: ** and * indicate significance at the 1% and 5% levels, respectively based on standard inference. p-values are in square brackets[.].							

In summation the hybrid models for the GBP/USD and the CAD/USD exchange rates fall short as many model coefficients are of the wrong sign so prompt a depreciation of the domestic currency as compared to a theoretical appreciation. The coefficient on the relative money supply in both models is far from unity so a one for one depreciation of the nominal exchange rate does not happen this result is akin to that of Lizardo and Mollick (2010) and Chinn and Moore (2011) for a model of the USD/JPY. A shortcoming of the hybrid model for the GBP/USD exchange rate is that the coefficients on three variables are insignificant namely the differentials related to income, the long-run interest rate and government spending (see Table 2.15.). Furthermore, tests of LE confirm that the relative income and relative long-run interest rate can be excluded from the model (see Table 2.13. Panel A.).

2.5. Conclusions

In this chapter, we review the USD vis-à-vis the GBP and CAD by application of the standard RID model (Frankel, 1979a) and a modified version of the standard model. The modified version (Hunter and Ali, 2014) is developed by inclusion of variables that have an impact on stable money demand and exchange rate persistence. The RID model is augmented by real share prices that account for long-run expectations as well as variables relating to productivity, government spending and the real oil price that account for real economic activity and external shocks to the exchange rate. We employ quarterly data from 1980:Q1 to 2015:Q4 a period marked by flexible exchange rate regimes and increased capital flows across borders. The RID model and modified version for the GBP/USD and CAD/USD are estimated by application of the Johansen (1995) methodology. We identify a single long-run cointegrating relation for both models but by tests of LE and WE we observe that the modified version yields a more coherent description of the identified long-run relation for the GBP/USD and CAD/USD. We observe that the modified model outperforms the standard RID model as the long-run underlying fundamentals that drive exchange rates and promote the breakdown of stable money demand are accounted for by the inclusion of real stock prices that react to one another in financial markets (Friedman, 1988). A more coherent presentation is in addition ascribed to the inclusion of variables that consider cross-border productivity differentials and the impact of non-traded goods on government spending. Another factor effecting performance is the inclusion of a real oil price that accounts for terms of trade and external shocks that bear impact on exchange rate persistence. The real oil price is however found to be weakly exogenous to the GBP/USD system so exhibits an indirect effect on the exchange rate whereas the real oil price is endogenous to the CAD/USD system as Canada is a major producer of oil, most of which is exported to the United States. We find that the nominal CAD/USD exchange rate has a reversed long-run projection so is driven by the system in the modified version as compared to forcing the system as was evident from individual WE tests for the RID model. We therefore deduce that parameters relating to real economic dynamics and the global financial architecture are central to models of long-run exchange rate determination.

We contribute to existing studies by refining the long-run projection of the standard RID monetary model for the GBP and CAD vis-à-vis the USD in a period characterized by flexible exchange rate regimes and financial market upheaval. It is evident from the hybrid monetary model that differentials linked to the real economy influence stable money demand and exchange rate persistence. We find that PPP is no longer a key driver of the exchange rate as financial assets play an important role in driving the system and micro and macro-economic phenomena cannot be characterised by short and long interest rates in a globalised financial market. A limitation of our work is that the VAR analysis employed is only as credible as its' identification schemes. We have however, addressed exogeneity in a framework otherwise characterized by endogeneity and omitted variables bias. The VAR analysis is in addition inadequate when it relates to structural inference and policy analysis (Stock and Watson, 2001). Our study can be extended by constructing a dynamic error correction model to investigate the short-run dynamics between model variables. Another recommendation for future research is to examine exchange rate determination and productivity differentials across diverse sectors (De Gregorio and Wolf, 1994) in the aftermath of the subprime mortgage crisis of 2008.

Appendix A2.1.

Table A2.1. Misspecification Tests of the RID Model for the Nominal GBP/USD Exchange Rate						
	Portmanteau(12)	LM(8)	ARCH(8)	Normality	Skewness	Excess Kurtosis
Panel A: Single-equation diagnostics using reduced-form residuals for the RID UK Model						
e*	9.15[0.33]	0.64[0.74]	0.78[0.62]	1.87[0.39]	-0.25	3.25
m-m*	10.34[0.24]	0.79[0.62]	1.2[0.31]	2.69[0.26]	-0.12	3.77
y-y*	15.06[0.06]	2.00[0.05]	1.00[0.44]	1.98[0.37]	-0.29	2.91
$i^s - i^{s*}$	17.19[0.03]*	1.46[0.18]	0.85[0.56]	2.23[0.33]	-0.12	3.2
$il - il^*$	14.5 [0.07]	1.77[0.09]	1.79[0.09]	4.82[0.09]	0.08	3.38
Panel B: System Tests LM (8) test: 1.22[0.06] Normality: 12.94[0.23] Notes: LM (8) Is a Lagrange multiplier test of serial correlation up to order 8; p values are reported in square bracketss [.].						

Table A2.2. Misspecification Tests of the RID Model for the Nominal CAD/USD Exchange Rate						
	Portmanteau(12)	LM(8)	ARCH(8)	Normality	Skewness	Excess Kurtosis
Panel A: Single-equation diagnostics using reduced-form residuals for the RID Canadian Model						
e*	9.11[0.33]	0.87[0.54]	2.51[0.01]*	7.77[0.02]*	0.19	3.57
m-m*	9.43[0.31]	0.93[0.49]	0.80 [0.59]	6.72[0.03]*	0.29	3.73
y-y*	9.30[0.32]	1.85[0.07]	1.32 [0.23]	1.27 [0.52]	-0.20	3.31
$i^s - i^{s*}$	11.63[0.16]	0.95[0.47]	2.57[0.01]*	2.06 [0.35]	-0.10	3.54
$il - il^*$	18.49[0.01]*	1.56[0.14]	0.89 [0.51]	0.68 [0.70]	-0.00	3.12
Panel B: System Tests LM (8) test: 1.38 [0.00]** Normality: 14.57 [0.14] Notes: LM (8) Is a Lagrange multiplier test of serial correlation up to order 8; p values are reported in square brackets [.].						

Table A2.3. Misspecification Tests of the Hybrid Model for the GBP/USD Exchange Rate						
	Portmanteau(12)	LM(1-8)	ARCH(8)	Normality	Skewness	Ex. Kurtosis
Panel A: Single equation tests						
e*	9.03[0.3]	1.13[0.35]	0.81[0.59]	1.77[0.41]	-0.12	3.07
m-m*	15.84[0.04]*	1.31[0.24]	0.51[0.84]	1.61[0.44]	0.00	3.28
y-y*	7.05[0.53]	1.47[0.18]	0.3[0.91]	0.83[0.65]	-0.1	3.22
$i^s - i^{s*}$	19.24[0.01]*	2.03[0.05]	1.1[0.32]	2.83[0.24]	-0.10	3.50
$i^l - i^{l*}$	19.83[0.01]*	2.22[0.03]*	1.44[0.18]	3.06[0.21]	-0.0	3.45
$s\text{-}s^*$	11.62[0.16]	1.14[0.3]	1.32[0.2]	1.14[0.56]	-0.22	3.23
$gs\text{-}gs^*$	7.06[0.52]	0.50[0.84]	0.50[0.84]	1.80[0.40]	-0.02	3.57
$prod^T\text{-}prod^{T*}$	12.01[0.15]	0.79[0.60]	1.00[0.43]	2.62[0.26]	0.05	3.47
roil	16.86[0.03]*	1.36[0.22]	1.02[0.41]	5.95[0.05]	-0.28	3.71
Panel B: System Tests						
LM(8): 1.21 [0.19] Normality: 22.18 [0.22] Notes: LM (1-8) Is a Lagrange multiplier test of serial correlation up to order 8; p values are reported in square brackets [.].						

Table A2.4. Misspecification Tests of the Hybrid Model for the CAD/USD Exchange Rate						
	Portmanteau(12)	LM(8)	ARCH(8)	Normality	Skewness	Ex. Kurtosis
Panel A: Single equation tests						
e*	20.48[0.00]**	2.03[0.05]	1.89[0.06]	1.44[0.48]	0.32	3.04
$m - m^*$	14.62[0.06]	2.55[0.01]*	1.44[0.18]	0.02[0.98]	-0.26	3.14
$y - y^*$	10.77[0.21]	2.05[0.05]	1.25[0.27]	0.83[0.65]	0.09	2.85
$i^s - i^{s*}$	21.82[0.00]**	2.69[0.01]*	0.44[0.89]	6.24[0.04]*	-0.28	3.93
$i^l - i^{l*}$	11.44[0.17]	1.43[0.19]	0.71[0.67]	7.21[0.02]*	-0.20	4.13
$s\text{-}s^*$	6.35[0.60]	0.31[0.95]	0.93[0.48]	2.41 [0.29]	-0.24	3.57
$prod^T\text{-}prod^{T*}$	19.78[0.01]*	0.72[0.66]	0.58[0.78]	1.27[0.52]	0.0	3.12
$gs\text{-}gs^*$	8.72[0.36]	0.77[0.62]	0.84[0.56]	4.08[0.12]	-0.42	3.59
roil	11.91[0.15]	1.45[0.18]	1.53[0.15]	1.34[0.22]	-0.34	3.27
Panel B. System Tests						
LM (8): 1.44 [0.02]* Normality: 32.99 [0.01]* Notes: LM (8) Is a Lagrange multiplier test of serial correlation up to order 8; p-values are reported in square brackets [.].						

CHAPTER THREE
THE PRICING RELATIONSHIP BETWEEN THE SPOT RATE AND FUTURES CONTRACT IN THE CURRENCY MARKET

3.1. Introduction

Limits to speculation render the assumption of risk neutral and rational market participants, invalid (Frankel and Froot, 1987; Froot and Frankel, 1989; Mussa, 1990). It is to be expected that spot exchange rates are affected by the behaviour of participants in the spot as well as the futures market for foreign currency since spot traders are likely to be influenced by information and expectations about the future.

Currency futures were first traded on the Chicago Mercantile Exchange (CME) in 1972. The currency futures market gained momentum following abandonment of the Bretton Woods regime in the 1970's and has evolved ever since. According to the Bank of International Settlements (BIS) 2016 Triennial Survey, trading in foreign exchange markets averaged $5.1 trillion per day in April 2016. Expansion of the currency futures market introduces another agency for systematic risk and cause for greater regulation in the derivatives market. Motive for examining the relationship between spot rates and futures contracts is to gain greater understanding of market functions and further transparency to assist regulators and policy makers in the establishment of appropriate industry directives.

Despite expansion of the foreign currency futures market there is a dearth of literature concerning underlying functions and efficiency. In section 3.2. we present a review of the relevant literature. In section 3.3. we outline a framework for modelling the pricing relationship related to the lead-lag relationship and market efficiency. In section 3.4. we model the pricing relationship between spot rates and futures prices. In section 3.5. we describe the data and econometric approach. In section 3.6 we present our empirical results. In section 3.7. we present concluding remarks and practical advice for future study.

3.2. Literature Review

In this section, we discuss literature related to the market for currency futures. It is reasonable to assume that due to the comparatively small size of the futures market that futures do not lead spot exchange rates (Dumas, 1996 and Lyons, 2001 pg. 113). Following the BIS 2016 Triennial Survey daily spot transactions for foreign exchange averaged 1.7 trillion USDs as compared to a daily average of 115 billion USDs for exchange traded derivatives including futures. Our study is pertinent in view of the growing market for foreign currencies. Theoretically, in a perfectly efficient market futures prices converge with spot rates through covered interest parity so that little information is assimilated in prices (Rosenberg and Traub, 2009). A discussion relating to the nature of futures contracts, futures and market completeness, the market for currency futures and the economic benefits of foreign exchange futures can be found in Appendix A3.1.

Recent literature on price discovery in the futures market encompasses currency futures and concludes, as is the case for stock index futures, that futures lead the spot market. It is therefore thought that the market for futures has an asymmetric impact on the spot market owing to the instantaneous flow of cost-free information between a high volume of diverse users in a concentrated space where a counterparty ensures swift transactions at a minimum cost (Huang, 2002 and Hasbrouk, 2003). The advent of revolutionary technology has accommodated expansion of the cash market for foreign currency and has perhaps given the cash market for currency precedence in price discovery. Several studies address the rate of price adjustment and direction of causality in the futures and cash market for foreign currency (Jochum and Kodres, 1998; Evans and Lyons, 2002; Wang et al., 2007; Martinez and Tse, 2008; Bekiros and Marcellino, 2013).

Martens et al. (1998) explore the mechanics of divergent arbitrage strategies and the effect of mispricing errors on index futures returns and index spot price returns by merger of a threshold autoregression model with an error-correction model and show that the greater the effect of mispricing error the greater the mispricing error itself. Martens et al. (1998) in addition show that the information impact of lagged S&P 500 futures returns on the corresponding price returns is particularly larger if the mispricing error is smaller than zero so that the return on index futures adjusts faster than the return on the spot price. Kurov and Lasser (2004) examine the link between E-mini S&P 500 and Nasdaq-100 futures traded on GLOBEX by exchange locals and off-exchange traders and show that traders near the trading pit exert a greater and prolonged impact on the price of E-minis owing to the instantaneous relay of information from the pit to the electronic trading system that allows for seamlessly swift transactions.

Tse et al. (2006) observe price discovery in two of the 'majors' or most traded currency pairs namely the USD/Euro and the JPY/USD and find that CME floor futures have little effect on spot rates while GLOBEX futures contracts exert the most influence on the spot Euro rate whereas retail online trading has the most impact on the spot JPY rate. Rosenberg and Traub (2009) investigate the relationship between spot rates and futures prices for the DMK, JPY, CHF and GBP against the USD to find that futures prices lead spot rates for the GBP and CHF by an interval of up to 3 minutes and by application of the methodologies developed by Hasbrouck (1995) and Gonzalo and Granger (1995) recognize futures information shares averaging between 80% and 90%.

Cabrera et al. (2009) study intraday data for the EUR and JPY vis-à-vis the USD on the CME GLOBEX market, E-mini futures and Electronic Broking Services over a four-month interval in 2005 and find by application of an information share approach, a common-factor component weight approach and an error correction adjustment approach that the spot market dominates price discovery so that spot rates lead futures prices. Chen and Gau (2010) in addition study the EUR-USD and JPY-USD around the time of macroeconomic announcements over a two-year period from January 2004 to December 2005 and show that the spot market is more informative to prices and so leads the futures market. Poskitt (2010) examines the role of the Reuters D3OO electronic dealer broker and the GLOBEX futures market in price discovery for the GBP/USD and shows that GLOBEX explains between 46 % and 47% of price discovery for the GBP/USD basis on regular assessments. Tornell and Yuan (2012) calculate econometric measures and build an algorithm as well as an investor sentiment index to delineate a relationship between spot rates and net positions in futures contracts for the GBP, EUR, JPY and MXN (Mexican Peso). Tornell and Yuan (2012) resolve that futures positions are only informative to the Euro over brief timespans despite the observation that peaks of speculative positions and troughs of hedging positions shed light on the price of futures.

Few studies address market currencies most likely due to a small market share and a dearth of coherent data. A pioneer study by Jochum and Kondres (1998) examines the impact of the futures market for the MXN, BRL and HUF (Hungarian Forint) vis-a -vis the USD on respective spot rates from January 1995 to February 1997 and by employing a Markov Switching Autoregressive Conditional Heteroscedastisity (MSWARCH) methodology as well as variance decomposition show that in the long-run spot rate volatility is tied to changes in the spot market rather than the futures market. De Boyrie et al. (2012) examine the relationship between the price of futures and spot rates for the BRL, ZAR and RUB by application of methodologies including a VECM, an open-end multiple structural analysis and a reduced form computation of the information shares to conclude that future prices lead the BRL spot rate and vice versa for the RUB but fail to define the lead-lag relationship for the ZAR. Another emerging market currency to be recently considered is the Indian Rupee (INR) most likely due to an increasing market share. Kumar and Trück (2014) examine the link between futures prices and spot rates for the INR/USD from September 2008 to January 2008 and by application of ADF tests reject unbiasedness for the futures premium as a predictor of variations in the spot exchange rate for contracts of longer maturities. Kumar and Trück (2014) in addition discern that spot currency returns, the futures basis and measures for realized volatility, kurtosis and skewness affect risk premiums at longer horizons.

In this section, we presented a review of the extant literature. The literature demonstrates that futures prices embody pertinent information about spot rates and that the futures market enhances market efficiency by balancing the cost of carry and reducing systematic risk. We find that in some instances the futures market has a disproportionate influence on the cash market and so in many cases leads spot rates. We observe from the literature that most studies pertain to the 'majors' or currencies with a large market share and that there is scope for research relating to emerging market currencies.

Since emerging market currencies are rarely explored we undertake to study the BRL and the ZAR (section 3.4. and 3.5.). Our study is significant considering the growing share of foreign exchange market transactions following the inception of the electronic trading platform in 1992. We undertake to examine 5-bilateral currency pairs in a period characterized by high volatility to gain greater understanding of causality in the market for foreign currency and to facilitate the need for transparency in the global financial market fraught by systematic risk.

3.3. The Framework for Modelling the Pricing Relationship

3.3.1. Testing the Lead-Lag Relationship

The lead-lag relationship may be examined by implementation of a type of Granger (1969) – Sims (1972) causality test as follows:

$$\Delta S_T = \alpha_0 + \alpha_I \sum_{i=-k}^{i=+k} \Delta F_{t-i} + u_t \tag{3.1}$$

where ΔS_t is the change in the spot price, ΔF_t is the change in the futures price and u_t is a white noise error term. The lead-lag relationship is determined by the significance of the coefficient on the change in futures price. The assumption is that if the lead coefficient is zero then the lag coefficient must be significant so by inference futures prices lead spot prices or vice a versa. However, in the instance of a feedback relationship the lead and lag coefficients tend to be different from zero. The Sims-Granger causality test cannot however supply information regarding the type and configuration of an asymmetric feedback relationship. Since spot and futures prices are attained concurrently (Stein, 1961; Anderson and Danthine, 1981; Garbade and Silber, 1983) WE is rejected but may still be examined. Stoll and Whaley (1990) are criticized for their application of OLS as Granger noncausality tests are only significant upon the assumption of WE. Chan (1992) acknowledges that prices are concurrently conditioned so applies the Generalized Methods of Moments (GMM) (Hansen, 1982) to address serial correlation and heteroscedastisity but neglects to adequately specify the model. We understand that equation (3.1) is misspecified owing to invalid conditioning as the system is not fully elaborated (lagged variables are excluded) and zero restrictions on ΔS_{t-1} and ΔF_{t-1}, i=1…k are not examined thereby suggesting the exclusion of possible cointegrating vectors. Oversight of lagged variables and invalid zero restrictions demand that spot and futures returns be martingale difference processes so that the emerging system is under identified and coefficient meaning inconsonant.

Apt specification of the pricing model considers the form of interaction between prices. Two frameworks exist for modelling the pricing relationship. The first framework examines departures of the futures price from a hypothesized value (MacKinlay and Ramaswamy, 1988; Yadav and Pope, 1990; Chung, 1991) while the second framework examines the lead-lag relationship between future prices and spot rates (Kawaller et al.1987; Harris, 1989; Stoll and Whaley, 1990). The first framework however imparts a deeper understanding of the nature of the lead-lag relationship observed in the second framework.

Cornell and French (1983a and b) construct the following model to describe the relationship between stock index futures and stock index prices:

$$F_{t,T}^{*} = S_t e^{r(T-t)} - \sum_{k=t+1}^{T} D_k e^{r(T-k)} \tag{3.2}$$

where $F_{t,T}^{*}$ is the theoretical stock index futures price quoted at time t for delivery at time T, S_t is the spot price of the stock index, r is the risk-free interest rate for the estimated time to maturity of the futures contract and D is the daily price return of the stock index for the time to maturity of the futures contract.

MacKinlay and Ramaswamy (1988), develop another framework estimated as follows:

$$F_{t,T}^{*} = S_t e^{(r-d)(T-t)} \tag{3.3}$$

where $F_{t,T}^{*}$ and S_t are described as above, r is the risk free interest rate, d is the income return on dividends from the underlying security portfolio and (r—d)(T—t) is a function for the cost of carry of the cash asset until expiration of the contract where (T-t) is the time to expiration. Pricing models however, examine departures of the realized futures price ($F_{t,T}$) from a theoretical value ($F_{t,T}^{*}$) quoted at time t for delivery at time T.

From Eq. (3.3) the function of the basis can be expressed as: $F_{t,T} - F_{t,T}^{*}$. The basis is an indicator of cost-of-carry arbitrage so that when a futures contract is overpriced or underpriced arbitrageurs will step in and sell or purchase the underlying asset thereby realigning prices to parity. Therefore, the basis marries the spot and futures market and identifies opportunities for arbitrage.

We describe the function of the basis as follows:

$$f_{t,T}^{*} = s_t + (r-d)(T-t) \tag{3.4}$$

where lower case letters indicate, the variables described in Eq. (3.3), in logarithmic form. If we assume perfectly efficient markets the realized price of the futures contract will be equal to the hypothesized value:

$$f_{t,T} - f_{t,T}^{*} = 0 \tag{3.5}$$

By substitution of Eq. (3.4) into Eq. (3.5) and repositioning we can recognize the relationship between the futures price and spot rate:

$$f_{t,T} - s_t = (r-d)(T-t) \tag{3.6}$$

We deduce from Eq. (3.6) that the basis will be equal to zero when the contract reaches maturity, since t will equal T, while for the duration of the contract we assume that the basis is equal to transaction fees but as the contracts nears expiration those fees become negligible.

3.3.2. Revision of the Conventional Approach to Lead-Lag Testing

As discussed earlier the model expressed in Eq. (3.1) is misspecified and in this section, we address invalid conditioning as well as the assumption of WE and the need to test for the assumption. We can assume Eq. (3.1) to be a VAR and since futures and spot rates are attained concomitantly, they are assumed to be endogenous to the system. Therefore, the lead-lag relationship is defined by a closed configuration as compared to a singular causal relationship.

Following Clements and Mizon (1991)[15] the lead-lag relationship is tested by testing zero restrictions in turn to attain a more compact specification of the VAR. The VAR once soundly specified can be incorporated in a structural model to be studied with tests for overidentifying restrictions. Let us consider the following closed system where $\mathbf{y}_t$ is an Nx1 vector of endogenous variables:

$$\mathrm{y}_t = \mu + \Pi_1 \mathrm{y}_{t-1} + \Pi_2 \mathrm{y}_{t-2} + \ldots\ldots + \Pi_p \mathrm{y}_{t-p} + \mu_t \tag{3.7}$$

where $\mathbf{\Pi}_i$ represents an NxN coefficient matrices, μ is a vector of deterministic components with μ_t as a white noise error term of mean zero and **Σ** denotes the variance-covariance matrix. By employing lag operator notation (3.7) can be expressed as:

$$\Pi(L) y_t = \mu + \mu_t \tag{3.8}$$

where (L) represents the lag operator and **Π**(L) denotes a p^{th} order matrix polynomial with $\mathbf{\Pi}_i = I_n$ expressing a NxN identity matrix.

The lead-lag relationship may in addition be expressed with variables in first differences as follows:

$$A(L)\Delta y_t = \mu + \mu_t^* \tag{3.9}$$

[1] Monfort and Rabemananjara (1990) discuss similar proposals for VAR models in stationary variables.

where Δ is the operator for first differences. Eq. (3.9) is the ‘non-structural’ representation of the examination of the lead-lag relationship. First differences are estimated to yield stationary series while the natural logarithm of a differenced series or variable yields the series’ returns. Furthermore, Engle and Granger (1987, pp.83-84) assume that:

1. In the absence of a deterministic trend a stationary series, denoted as a nonsingular autoregressive moving average process that has been differenced is described as integrated of order d represented as $y \sim I(d)$.[16]
2. Elements of the vector y_t are described as being cointegrated of order d,b, represented as $y_t \sim CI(d,b)$ so that when every element of y_t is integrated of order I(d) a cointegrating vector α (≠0) comes into being so that $z_t = \alpha' y_t \sim I(d-b)$, b>0.

The first assumption accommodates Eq. (3.9) that will henceforth be described as the 'quasi-unrestricted' model. The ‘quasi unrestricted’ model, that is first differenced, neglects to impose zero restrictions on the coefficients of lagged dependent variables despite the placement of common unit root restrictions. The 'traditional' Granger causality test tends to overlook the second assumption that together with unsound conditioning gives rise to misspecification.

Regarding the Granger Representation Theorem (Engle and Granger, 1987), in the instance of cointegration where d=b=1 and both assumptions above hold, we can extract an error correction model of the VAR, granted that cointegration and error correction models hold a similar form. When the futures price and spot rate are nonstationary, or integrated of order 1, we presume that both variables when first differenced will be stationary or I(1). It is in addition plausible that from assumption (2) both variables cointegrate so that a linear

[16] That is, for some p and q y_t will belong to the ARIMA (p,d,q) class of models proposed by Box and Jenkins (1970).

combination of the spot and futures prices, that are individually integrated of order 1, will be stationary or I(1). This conclusion is observed from the mispricing literature where the basis is assumed to be the error correction mechanism and the cointegrating vector equal to unity.

We conclude therefore that the frameworks set out in Eq. (3.1) and Eq. (3.9) will be misspecified except where there is no cointegration and the white noise error term is insignificant in all the system equations. Bypassing the theoretical considerations of long-run parity in the computation of the first differenced VAR leads to misspecification and false inferences about the nature of the lead-lag relation since tests of zero restrictions are false.

The same holds for estimation of Eq (3.9) in a moving average framework to observe impulse response functions. We therefore consider employing Eq. (3.8) in terms of the error correction (Johansen, 1988 and Johansen and Juselius, 1990) as specified below:

$$\Delta y_t = \mu + \Gamma_1 \Delta y_{t-1} + \ldots\ldots + \Gamma_{k-1} \Delta y_{t-k+1} + \Pi y_{t-k} + \mu_t \qquad (3.10)$$

where the expression for lagged variables y_{t-1} denotes the equilibrium error in the short-term, or departures from the long-term parity, and of which a minimum of a single component must be non-zero. Π supplies information regarding the effect of the equilibrium error. By following Johansen (1988) and Johansen and Juselius (1990) the long-term response matrix is inserted in $\Pi=\boldsymbol{\alpha\beta}'$ where $\boldsymbol{\beta}'$ is the matrix of cointegrating coefficients such that $\boldsymbol{\beta}' y_{t-1} \sim \mathrm{I}(1)$ and $\boldsymbol{\alpha}$ is the matrix of adjustment coefficients. As inferred by Johansen (1995) the matrix Π will exhibit reduced rank if any of the variables are cointegrated in the vector y and so the foundation for the Johansen (1988) test of cointegration is outlined. Rank (π) is represented by r and three prospects exist:

1. r=0, in this instance all of the variables are I(0) and there is no evidence of cointegrating vectors.

2. r=N in this instance every variable is I(1) and there are N cointegrating vectors as all linear combinations comprising stationary variables are stationary.
3. 0<r<N in this instance r linear combinations exist comprising the nonstationary variables that are stationary. Likewise, N-r common stochastic trends will be present (Sims et al., 1990).

Once mispricing and the lead-lag relationship are investigated jointly and contingent upon each other the matter of effectively functioning markets can be addressed by a single notion. The single notion of mispricing and the lead-lag relationship can be expressed as a reduced form VAR following the notation:

$$Y = XB + e \tag{3.11}$$

Ordinary Least Squares can be employed to estimate Eq. (3.11) as it is a VAR, the system estimator is expressed as follows:

$$\hat{B} = [I \otimes (X'X)^{-1} X']Y \tag{3.12}$$

As argued previously the futures price will equal that of the spot at maturity of the futures contract so by means of short-run departures from long-run parity the spot and futures prices will intersect as maturity nears. The inference is that as the time to expiration declines, (T-t) →0, X approaches Y, X→Y, and the estimator for B becomes I as follows:

$$\hat{B} = [I \otimes (Y'Y)^{-1} Y']Y \rightarrow \hat{B} = I \tag{3.13}$$

By substitution in (3.11) we derive:

$$Y = X + e \rightarrow Y = Y \text{ as } \lim_{t \to T} X = Y \text{ and } e = 0 \tag{3.14}$$

So that X and Y cannot be distinguished from one another and an issue of identification arises from the long-run parity.

$$\begin{bmatrix} \Delta f_t \\ \Delta s_t \end{bmatrix} = \begin{bmatrix} \alpha_{10} \\ \alpha_{20} \end{bmatrix} + \begin{bmatrix} -\alpha_{10} \\ \alpha_{20} \end{bmatrix} \left[f_{t-1} - s_{t-1} \right] + \begin{bmatrix} u_{1t} \\ u_{2t} \end{bmatrix} \tag{3.15}$$

However, if the restricted VAR expressed in Eq. (3.15) above is an exact model then the system cannot be identified, no matter the period to maturity.

Therefore, when mispricing and the lead-lag relationship are treated as a single notion and conditional on one another the incapacity to identify the system suggests that the foreign currency market functions effectively the spot and futures price converge and become indistinct.

We conclude that issues of misspecification surface when prescience from the study of mispricing is disregarded upon examination the lead-lag relationship. Also, unsound results can emerge from neglecting to treat mispricing and the lead-lag relationship as a single structure as the models commonly employed cannot be identified. The estimation of GMM or the adoption of auxiliary variables to appropriately instrument does not however override the obstacle in this case while the application of OLS to a structural' equation produces biased and inconsistent estimates since any assumption of exogeneity is not only unsound but cannot be unexamined.

Mispricing studies indicate that homogeneity is a characteristic of the long-run parity between spot and futures prices such that short-run departures from this parity are produced by the basis. Garbade and Silber (1983) corroborate this result by application of a reduced form VAR in a study of spot and futures prices for commodities.

In union, mispricing and the lead lag relationship infer that the structure pertaining to spot and futures prices cannot be identified so that the two markets are one and the same. In this instance, foreign currency markets can be described as effectively functioning. The combined framework permits us to elaborate on the description of ‘effectively functioning'. Effectively functioning financial markets are classified as either strongly effectively functioning or weakly effectively functioning. If the markets do not correspond to either category, they are deemed to be ineffectively functioning. Markets need to comply with several conditions to meet either category. The conditions pertaining to the effectively functioning categories are outlined below.

1. Strongly effectively functioning
    a. The price series cointegrate with the cointegrating vector that is the basis. Thus, the homogeneity constraint must remain.
    b. System identification cannot be established from the reduced form.
    c. The pricing relationship does not alter as time evolves so that the reduced form remains steady.

    Under this set of conditions prices in the spot and futures market will be identical so that prices are contingent upon the basis or the extent of mispricing which is of consequence to arbitrage activity. If a single or two conditions are maintained, then they are necessary but not sufficient to guarantee strongly effectively functioning markets. If all the conditions are satisfied, then they are both necessary and sufficient to warrant strongly effectively functioning markets.

2. Weakly effectively functioning

    Conditions a. and c. need to be maintained to support the strongly effective functioning of a market. Regarding condition b. if individual equations in the reduced form system are identified, then the markets are deemed to be weakly effectively functioning. Once more, if a single or two conditions are maintained then they are necessary but not sufficient to guarantee strongly effectively functioning markets. If all the conditions are satisfied, then they are both necessary and sufficient to warrant strongly effectively functioning markets. If markets are weakly effectively functioning, then regulators can decipher the most appropriate directive to be implemented.

If markets are deemed to be ineffectively functioning, then the origin or a hint of the origin of ineffectiveness can be determined inside the joint framework explicitly when conditions b. and c. for strongly effectively functioning markets are maintained but cointegration is not established or the cointegrating vector is not the basis. This infers that the markets are ineffectively functioning and that the arbitrage process is ineffective. Thus, the trading systems may require amendment or establishment of a new directive.

If conditions a. and c. for strongly effectively functioning markets is maintained but identification of the system cannot be proved, then markets are ineffectively functioning as prices do not internalize the greatest amount of accessible information. An argument for this may be stickiness in the trading system. If conditions a. and b. for strong effectiveness are maintained but the system is unsteady the pricing relationship could have evolved over time which is evidence of ineffectively functioning financial markets. Once again, this information can be employed to determine the origin or hint at the origin of ineffectiveness within the joint framework. The capacity to categorize markets in this manner permits regulators to devise directives that are appropriate to individual markets.

3.3.3. Market Efficiency and Effective Functioning

The framework for examining effective functioning can be implemented, with minor adjustment to the explanation, to examine the efficiency of the spot and futures market in unison. Therefore, efficacy and efficiency become one and the same. We adapt the concept in consideration of market efficiency and present conditions for a system that demonstrates efficiency akin to those presented for effectively functioning markets but less restrained.

Distinct from the Fama's (1970) definition of market efficiency we present the definition as argued by Dwyer and Wallace (1992) which refers to the review of exchange rate market efficiency by Levich (1985) and that Ross (1987) considers the more useful definition. A market is described as efficient in the absence of opportunities for profit which would broaden the agent's expected utility. More specifically as defined by Dwyer and Wallace (1992):

'It is hard to see how a market with no expected utility increasing profit opportunities available to agents based on expected utility maximizing acquisition of information could be characterized as inefficient in any interesting sense of the word.'

Spot market efficiency has generally been examined with respect to the relationship between spot and forward prices. The most prominent procedure for examining spot market efficiency is to observe whether forward prices are an optimal predictor of future spot rates or if the forward rate is biased and has a risk premium, that may vary over time. The hypothesis of the forward rate being an optimal predictor of future spot rates has been investigated by a vast number of scholars including Hansen and Hodrick (1980 and 1983), Baillie, Lippens and McMahon (1983), Fama (1984), Domowitz and Hakkio (1985), Hodrick and Srivastava (1986), Backus et al. (1993), Baillie and Bollerslev (2000), Meredith and Ma (2002) and Frankel and Poonawala (2010) amongst many others.

Tests of the forward rate as an optimal predictor of the future spot rate are conducted under the assumptions of rational expectations and risk-free arbitrage and are expressed as follows:

$$E(S_{t+n} | \Omega_t) = F_{t,t+n} \tag{3.16}$$

where S_{t+n} is the spot rate at time t+n, $F_{t,t+n}$ is the forward rate fixed at time t for an underlying asset to be delivered at time t+1, Ω_t is the information set as of time t and E(.) is the mathematical expectations operator. The assumption of rational expectations suggests that

$$S_{t+n} = E\left(S_{t+n}\middle|\Omega_t\right) + \varepsilon_{t+n} \text{ where } E\left(\varepsilon_{t+n}\middle|\Omega_t\right) = 0 \quad (3.17)$$

where ε_{t+n} is a zero mean, moving average error process of (n-1)th order. By substitution of Eq. (3.16) in Eq. (3.17) we derive:

$$S_{t+n} = F_{t,t+n} + \varepsilon_{t+n} \quad (3.18)$$

Equation (3.18) asserts that the forward rate quoted at time t for delivery at t+n is an unbiased predictor of the spot rate at time t+n. In addition, ε_{t+n} now presents the forecast error. Eq. (3.18) can be examined by the representation:

$$S_{t+n} = \beta F_{t,t+n} + \varepsilon_{t+n} \quad (3.19)$$

where the condition for efficiency is that β=1.

Examination of efficiency founded on the forward rate as an optimal predictor of the future spot rate has generated ambivalent results. When Eq. (3.25) is applied to examine market efficiency the condition β=1 is inclined to be accepted and the forward rate is deemed to be an unbiased and therefore optimal predictor of the future spot rate. The model in Eq. (3.19) has in addition been specified as follows:

$$\Delta s_{t+n} = \alpha + \beta(f - s)_t + \varepsilon_{t+n} \quad (3.20)$$

where lower case letters represent the variables in logarithmic form.

Tests of efficiency that apply the specification in Eq. (3.20) have rejected the restrictions α=0 and β=1 showing evidence of a risk premium and presenting the forward rate as a biased predictor of the spot rate risk premium (α≠0). These ambiguous results arise from the applied approach. $\beta=1$ is generally not rejected based on tests applying the model in Eq. (3.19) since the two prices almost exactly trail each other. The model therefore misinterprets the relationship between the two prices as the trend that they both closely follow. The latter acknowledgement is detrimental to examination of the relationship as the series are nonstationary so standard inference procedures in this instance are invalid.

The model in Eq. (3.20) has been rejected by some (Hakkio, 1981) due to model misspecification. Hakkio and Rush (1989) acknowledge that spot and futures prices are stochastic nonstationary and can be perceived as a cointegrating regression so construct a framework based on this feature to address the predicament previously discussed.

A condition for cointegration is stationarity so to begin one tests for a unit root in the basis (s_{t+1}-$f_{t,t+1}$). Therefore, cointegration where $\beta=1$ is a condition for efficiency. Dwyer and Wallace (1990) highlight that studies showing (s_{t+1} -$f_{t,t+1}$) is the cointegrating vector are incorrect with the interpretation that cointegration validates inefficiency (Baillie and Bollerslev, 1989)[17] as lack of cointegration is confirmation of inefficiency.

If the forecast error is found to be the cointegrating vector then this is insufficient proof of efficiency. Subsequently the next stage is to examine efficiency in the reduced form error correction model that is accompanied by evidence of cointegration. By setting n=1, we estimate equation 3.21, below, to examine the joint hypothesis that H_0:-$\rho=\alpha=\beta=1$.

$$\Delta s_{t+1} = \alpha \Delta f_{t,t+1} + \rho(s_t - \beta f_{t-1,t+1}) + \varepsilon_{t+1} \tag{3.21}$$

If the hypothesis is not rejected, then Eq. (3.21) breaks down to Eq. (3.19) so that we obtain

[17] Baillie and Bollerslev (1989) view the markets for spot and forward foreign exchange as distinct speculative markets as they contend that the error correction term can forecast prices.

market efficiency. In addition, when the basis is stationary, the test statistics from (3.21) are more robust as standard inference can be applied (Kremers et al., 1992).

The notion is relevant to stock index markets, spot and futures, but requires some adjustment that renders moderate conditions for efficiency. The prices of futures contracts will be determined efficient by the futures market if the realized futures price is commensurate with:

$$f_{t,T}^{*} = s_t + (r-d)(T-t) \tag{3.22}$$

Where the long-run parity is f=s indicating the restriction β=1 in $f_t = \beta s_t + \varepsilon_t$. This can be examined by investigating the unit root properties of the basis. If the basis is shown to be stationary then the restriction β=1 is valid. The reduced form error correction model that ensues is represented as follows: $\begin{bmatrix} \Delta f_t \\ \Delta s_t \end{bmatrix} = \begin{bmatrix} -\alpha_{11} \\ \alpha_{21} \end{bmatrix} [f_{t-1} - s_{t-1}] + \begin{bmatrix} u_{1t} \\ u_{2t} \end{bmatrix}$ (3.23)

and by construction of the error correction term the system subsequently condenses to f=s (s=f). The upshot of this systems approach is that efficiency in both markets can be observed in unison.

The condensed framework calls for conditions that diverge from those already presented but are not very distinct from those stipulated for strongly effectively functioning markets. The conditions are outlined as follows:

1. Both markets are efficient

   If both markets are described as efficient then the basis must be the cointegrating vector and the entire system is not identified. In the case that both stipulations hold then in unison they are necessary and sufficient to establish market efficiency. The contrast between efficient and strongly effective is that no rationale exists for demanding stability but if stability is demonstrated then the market is thought to be perpetually efficient. However, this is rarely the case especially in the instance of stock markets. If the model is unstable then the market will be efficient for the duration of the sample.

2. One market is efficient

   If the basis is the cointegrating vector and one of the equations is under identified then the equation that is identified is commensurate with the market that is inefficient. An inefficient market would indicate opportunities for arbitrage in the presence of excess returns, relating to lagged returns in either market or both, in the equation such that these returns can indeed be predicted with regard to the market identified as inefficient. Stability in this instance is not a stipulation.

3. Both markets are inefficient

   In the absence of cointegration or when the basis is not found to be the cointegrating vector then both markets can be described as inefficient. In this instance, the spot and futures market will not shadow each other and as the distance between the two grows so too does the creation of excess returns thus enabling market participants to formulate a plan that takes advantage of market inefficiency. Where the basis is not the cointegrating vector but there is evidence of cointegration prices can be forecast so that market participants can capitalize on atypical returns. In the last instance where both equations are identified the inference is that market participants can once again predict future returns and devise a plan to profit from anomalies.

The incongruity with this notion, as is with the efficient-market hypothesis, is that if markets are efficient hedgers, speculators and arbitragers (mostly arbitragers) enter to exploit the presence of excess returns so that prices return to parity. Where markets are perceived to be efficient, but are not, participants adopt passive investment strategies and due to sparse transactions inefficiencies arise. Thus, market prices are not realigned so that they are accurate and fully reflect all available information.

3.3.4. Is Mispricing Path Dependent?

In this section, we discuss ramifications of the previous discourse for the behaviour of mispricing. We contemplate whether the model presented in section 3.2.3 can expand upon path trajectory and provide insight as to whether mispricing is path independent or path dependent. We claim that the logic behind the model's formulation has direct ramifications for the mispricing path. MacKinlay and Ramaswamy (1988) and Yadav and Pope (1990) corroborate path dependence in a gauge for respectively mispricing the S&P 500 Index futures and the FTSE 100 Index futures. Brennan and Schwartz (1990) develop a framework for modelling the properties of mispricing that recognizes the progressive properties of mispricing that follow a continuous time stochastic progress noted as a Brownian Bridge process and expressed as follows:

$$d\varepsilon(\tau) = -\frac{\mu\varepsilon}{\tau}dt + \gamma dz \tag{3.24}$$

where ε is mispricing, τ is the interval to expiration of the futures contract and μ is the rate of reversion to mean. This stochastic process is distinct as it exhibits path independence when mispricing relapses to zero and is commensurate to zero with a probability of one when τ=T or at the time of contract maturity.

If we contemplate features of mispricing (the basis) and the examination of path dependence and consider that conventional tests of the lead-lag relationships are misspecified as they disregard cointegrating vectors. A cointegrating vector will emerge by cointegration if a linear combination of two stochastic nonstationary, I(1), series is stationary, I(0). External shocks to a stochastic nonstationary series are permanent as compared to transient as is the case for a deterministic nonstationary series. It is therefore illogical to contemplate that the cointegrating vector which is a linear combination of the two nonstationary series will be cointegrated or integrated of order zero and follow an autoregressive moving average process of order (p,q).

The basis or cointegrating vector is therefore a measure of mispricing and is a stationary stochastic process according to the prior argument. It is possible that the two series are driven by the non-stochastic trend that may then be common. Path dependence in mispricing is examined by regressing the basis on the time to maturity and observing if time to maturity is important. The interval to maturity is given by (T-t) where T is the time to contract maturity. The coefficient sign subsequently determines the type of the path dependence. Time to maturity is measured by (T-t) where T is the maturity date, set at the onset of the contract. t is the amount of trading days remaining during the contract. (T-t) therefore denotes the number of trading days to contract expiration. As T is set at the onset of the contract, (T-t) diminishes by one unit every day until maturity. (T-t) can therefore be described as a deterministic time trend. When mispricing is determined to be stationary around a deterministic trend it is considered significant, but this negates the definition of cointegration. So, if the prices of spot and futures are stochastic nonstationary and the basis is the cointegrating vector then mispricing is stochastic stationary, which is consistent with the above.

In view of this consideration the evidence of path dependence presented by MacKinlay and Ramaswamy (1988) and Yadav and Pope (1990) is nothing more than model misspecification. Yadav and Pope (1990) confirm model misspecification when they model mispricing and demonstrate that it is a 1st order autoregressive process. Furthermore, a path independent model that is mean reverting such as that developed by Brennan and Schwartz (1990) has proven to be most suitable for modelling the theoretical nature of mispricing.

In this section, we establish a framework to examine effective functioning of the stock market and stock index futures market which can be applied to other markets including the market for foreign exchange. In section 3.5. we empirically examine the theory that the foreign exchange futures market leads the underlying cash market and apply the framework discussed above to review the efficiency of the market for foreign exchange in entirety.

3.4. Modelling the Pricing Relationship

In this section, we investigate price discovery in the market for foreign exchange and apply the framework developed in section 3.2. We address points previously raised to gain an understanding of market functioning. If the foreign exchange market is determined to be strongly effectively functioning, then the futures market is executing its function effectively. In addition, if the markets are found to be strongly effectively functioning they are efficient and opportunities for arbitrage are non-existent. Weakly effectively functioning markets can in addition be described as efficient.

We also address the behavior of mispricing or the cointegrating vector, identified as the basis, which is an indicator of the degree of mispricing. The basis is assumed to be mean reverting or a stationary stochastic process this assumption follows that of Brennan and Schwartz (1990).

The points in question are examined by application of the model formulated by Brennan and Schwartz (1990) expressed as follows:

$$\begin{bmatrix} \Delta f_t \\ \Delta s_t \end{bmatrix} = \begin{bmatrix} a_{10} \\ a_{20} \end{bmatrix} + \begin{bmatrix} -a_{11} \\ a_{21} \end{bmatrix} \left[f_{t-1} - s_{t-1} \right] + \begin{bmatrix} u_{1t} \\ u_{2t} \end{bmatrix} \tag{3.25}$$

which is a 1st order reparameterised VAR represented in error correction form. Conditions for effectively functioning markets are derived from the above model and to reiterate for effectiveness and efficiency we call for:

1. Long-run homogeneity meaning that the error correction term is the basis.
2. Under identification of both equations contained in the system.

While for strongly effectively functioning markets but not efficient markets we call for:

3. Model stability

If all three conditions are maintained, then the spot and futures market are contingent on the basis.

In section 3.5. we describe the data and econometric approach and in section 3.6 we present our empirical results and examine effective functioning in the foreign exchange market and the rationale for the market efficiency conditions described in section 3.3.3. as well as the stochastic properties of the basis.

3.5. Data and The Econometric Approach

3.5.1. Data

Following our discussion in the previous section and the assumptions made by Engle and Granger (1987), we examine the stationarity properties of the individual variables in our sample and apply the Johansen (1995) methodology to assess cointegration with the basis. To test these conditions and model the pricing relationship we employ weekly observations of spot and futures exchange rate data for the Canadian Dollar (CAD), British Pound (GBP), Australian dollar (AUD), Brazilian Real (BRL) and South African Rand (ZAR) against the of the United States Dollar (USD).

We collect weekly data for spot exchange rates and futures contracts traded on the CME from Thompson Reuters Datastream. Futures contracts for the developed market currencies expire quarterly (March, June, September and December) while those for the emerging market currencies expire monthly. Data for the USD/AUD and USD/GBP exchange rate futures contracts is for the period 1994:Q1-2015:Q4. This period is chosen to account for variance in the USD/GBP following Britain's opt-out of the Exchange Rate Mechanism in September 1992 and enforcement of the Maastricht Treaty in November 1993. Data for the USD/CAD exchange rate contracts to compare with the USD/AUD and USD/GBP currency contracts is only available from 1995:Q1-2015:Q4. Our sample period for the USD/BRL and USD/ZAR currency contracts is for the period 1998:Q1-2015:Q4 and considers a period characterized by inflation targeting in both economies.

The USD/GBP exchange rate is examined as it is the world's third most traded currency pair and gained shares in world foreign exchange turnover averaging 9.3% daily in April 2016.[18] The USD/CAD and USD/AUD exchange rates are respectively the fourth and fifth most traded currency pairs with daily turnovers of 5.2% for the USD/CAD and 4.3% for the USD/AUD. Emerging market currencies are rarely explored perhaps due to a small market share and lack of coherent data. We undertake to study the currencies of Brazil and South Africa as these economies are comparable in their economic projection and have both endured periods of hyperinflation. In addition, Brazil as well as South Africa experienced periods of appreciation prior to de-valuation at the time of the 2008 financial crisis when

[18] The Bank of International Settlements (BIS) 2016 Triennial Survey
http://www.bis.org/publ/rpfx16fx.pdf

investors sought to invest in safe haven currencies as compared to the risky currencies of developing nations. The USD/BRL and USD/ZAR exchange rates are respectively the ninth and tenth most traded emerging market currency pairs with daily turnovers averaging 0.9% and 0.8% in April 2016.

Regarding inflation targeting in Brazil and South Africa. The period surveyed comes after South Africa's first democratic elections in 1994 and subsequent establishment of the Government of National Unity that was levied with the task of reducing domestic inflation. In a similar vein, the Central Bank of Brazil, to curb inflation, was pressed to float the Real in January 1999 following unsuccessful implementation of the Plano Real that prompted a period of hyperinflation and a devaluation of the national currency.

The futures price is constructed as a sequence rather than a series. We extract three months of futures prices quoted for each nearest maturity contract (January, March, September and December) for the USD vis-à-vis the AUD, CAD and GBP. While a month's futures prices are extracted for each nearest maturity contract (monthly) for the USD vis-à-vis BRL and ZAR. The spot rates are in addition extracted so that they are analogous with the futures prices.

By visual inspection of the graphs in Appendix A.3.2. (Figures A3.1.-A3.13.) we observe that the two prices almost completely shadow one another at 1, 2 and 3 months prior to expiry for the USD/GBP, USD/AUD and USD/CAD exchange rates and at 1 month and two weeks prior to expiry for the USD/BRL and USD/ZAR exchange rates. From Figures A3.1.-A3.13. we discern two significant events namely the collapse of the dot-com bubble in late 2002 and the collapse of the shadow banking system in September 2008. It is striking that the two prices continue to follow each other through periods of financial upheaval. A slight departure from the USD/BRL spot exchange rate is observed in Figure A3.10. for the futures price one month prior to expiry in Q4:2008. This deviation is most likely a consequence of investors' flight to safety and withdrawal from emerging markets at the onset of the 2008-2009 financial crisis. We in addition note that the two series appear to behave like a random walk. We continue to test whether the series are indeed first difference stationary.

3.5.2. Econometric Approach

In this section, we examine whether the pricing relationship varies over time. We recognize that transaction costs are negligible in the near period to maturity and at maturity itself so that the two prices are equal. If the prices do not match at expiry the markets are said to be inefficient and have therefore generated excess returns and worthwhile opportunities for arbitrage. We investigate whether the spot and futures price follow each other at longer periods prior to expiry. Since the futures contracts for the AUD, CAD and GBP are quarterly we analyze the pricing relationship one, two and three months prior to expiry. The contracts for the BRL and ZAR are monthly so we examine the pricing relationship at two weeks and one month prior to expiry.

Our analysis comprises tests of the univariate time series properties of the spot exchange rate and respective futures price for the individual currency contracts at different intervals to expiration. It is inferred that economic and financial time series have a unit root in their autoregressive representation and can be referred to as stochastic nonstationary (Nelson and Plosser, 1982; Perron, 1988; Baillie and Bollerslev, 1989). Stochastic nonstationary series yield a constant mean in their first difference so are dubbed difference stationary (Nelson and Plosser, 1982). Nonstationary series may in addition indicate a deterministic trend so that departures from the trend are stationary.

To distinguish between stochastic nonstationary and deterministic nonstationary we present the following models:

$$y_t = \alpha + \beta t + \mu_t = \rho\mu_{t-1} + e_t, e_t \sim N(0, \sigma_e^2), |\rho| < 1 \qquad (3.26)$$

where t represents the deterministic time trend (t=1,2, … T) and the errors pursue a first order autoregressive process. On the other hand, Eq. (3.27) represents a random walk with drift on condition that μ≠0 as follows:

$$y_t = \mu + y_{t-1} + v_t, v_t \sim N(0, \sigma_v^2) \qquad (3.27)$$

If we assume that y_t=μ, by reiterative substitution Eq. (3.2) can be expressed as:

$$y_t = t\mu + \sum_{i=1}^{t} v_i \qquad (3.28)$$

To distinguish between the two series, we take the s-period prior to the forecast of Eq. (3.26) and Eq. (3.27) and correspondingly derive:

$$E(y_{t+s} | y_t) = \alpha + \beta(T+s) + \mu_{t+s|t} = \rho^s \mu_t \qquad (3.29)$$

$$E(y_{t+s} | y_t) = s\mu + y_t \qquad (3.30)$$

Substituting Eq. (3.28) into Eq. (3.30) we obtain

$$E(y_{t+s} | y_t) = \mu(t+s) + \sum_{i+1}^{t} v_i \qquad (3.31)$$

In Eq. (3.29) as s approaches infinity or $\lim_{s\to\infty}$ we observe that $\rho^s\mu_t \to 0$ and the series becomes independent as it is freed of errors so that subsequent shocks are temporary. On the other hand, shocks are intransient and persist for the series with a stochastic trend. We can in addition show that the two series diverge by taking the first difference of Eq. (3.26) (assuming no serial correlation) and (3.27).

Another approach to recognize the first difference of (3.27) is to observe the unit root as a coefficient restriction. So, for Eq. (3.27) we derive:

$$\Delta y_t = \mu + v_t \tag{3.32}$$

which, is stationary by the definition of white noise.

We observe a difference in Eq. (3.26) when expressed as a coefficient restriction as follows:

$$\Delta y_t = \beta + \mu_t - \mu_{t-1} \tag{3.33}$$

We present the AR(1) error in Eq.(3.26) as:

$$\begin{aligned} \mu_t &= \frac{e_t}{1-\rho L} \\ \Rightarrow \Delta\mu_t &= \frac{e_t}{1-\rho L} e_t \end{aligned} \tag{3.34}$$

When ρ=0, then differencing the trend stationary series will create a moving average error with a singular unit root. An invertible unit root would cause non-stationarity in the error term and a spurious regression (Granger and Newbold, 1974).

Several methods have been developed to test the stochastic properties of a time series. Popular approaches for testing the unit root hypothesis or non-stationarity include those developed by Phillips-Perron (Phillips, 1987; Perron, 1988; Phillips and Perron, 1988), Dickey-Fuller (1979), Ellliot, Rothenburg and Stock (1996), Kwiatkowski-Phillips, Schmidt and Shin (1992) and Ng and Perron (2001). The stochastic properties of pooled series may in addition be tested by application of panel unit root tests. Methodologies for testing the unit root properties of pooled data include those of Im, Pesaran and Shin (1997 and 2003), Levin, Lin and Chu (2002), Breitung (2001) and Hadri (2000).

The Dickey-Fuller (1979) test is used widely but additional lags are required to fulfill the condition for residual serial correlation. If the time series is correlated, the assumption of white noise errors ε_t is rejected. A parametric correction is formulated by the ADF test for correlation of a higher order by inferring that the time series under investigation follows an autoregressive process of order p and by inclusion of p lags of the dependent variable. The Phillips-Perron (1988) test on the other hand is a nonparametric method that permits heterogeneity as well as serial correlation by computation of a consistent estimator of the variance in the unit root processes.

Examining the unit root process of a time series by application of the Phillips-Perron (1988) test comprises estimation of the following three models:

$$y_t = \mu + \beta(T - \frac{n}{2}) + \mu y_{t-1} + \mu_t \qquad (3.35a)$$

$$y_t = \mu^* + \alpha^* y_{t-1} + \mu_t^* \qquad (3.35b)$$

$$y_t = \hat{\alpha} y_{t-1} + \hat{\mu}_t \qquad (3.35c)$$

and by employing the modified F or t statistics (denoted as Z statistics by Phillips-Perron, 1988) to examine appropriateness of the null hypothesis.

In addition, by estimation of Eq. (3.35a) we can test three hypotheses:

1. H_0: $(\mu, 0, \alpha) = (\mu, 0, 1)$ tests for a stochastic trend with drift by employing the modified F statistic, $Z(\Phi_3)$, as compared to the alternative that the unit root process is deterministic nonstationary.

2. H_0: $(\mu, \beta, \alpha) = (0,0,1)$ tests for a stochastic trend without drift by employing the statistic $Z(\Phi_2)$.

3. H_0: $\alpha=1$ by employing the $z(\alpha)$ or $Z(t_\alpha)$ statistics.

For Eq. (3.35b), the null hypothesis is H_0: $(\mu^*, \alpha^*) = (0,1)$ or H_0: $\alpha^*=1$. The above hypotheses are examined by employing the $Z(\Phi_1)$ and $Z(\alpha^*)$ or $Z(t_{\alpha^*})$ statistics correspondingly. The opposing hypothesis to the above three is that the series is a stationary stochastic process and α*<1. For Eq. (3.35c) the null is $H_0 : \hat{\alpha} = 1$ and the alternative is α*<1, as for Eq. (3.35b) while the test statistic is either $Z(\hat{\alpha})$ or $Z(t_{\hat{\alpha}})$.

The reason for specifying three models to test the unit root hypothesis becomes clear when we consider the testing strategy to be adopted. Dickey, Bell and Miller (1986) maintain that the model in Eq. (3.35b) is the most appropriate for a stochastic nonstationary series with drift since it exhibits greater explanatory power than the model in Eq. (3.35a). The model in Eq. (3.35b) would suit a stochastic nonstationary series without drift as compared to model (3.35c) that exhibits low explanatory power for series with both trend and drift.

Perron (1988) demonstrates that it is difficult to discern between a deterministic trend and unit root in large samples however overlooking the presence of a trend would prove detrimental to the model. Perron (1988) therefore resolves to test for a unit root by estimation of the model with trend and drift and if the hypothesis is not rejected to examine the unit root properties by means of the model without drift. If the unit root hypothesis is rejected by the model without drift then the most suitable model is the one expressed in Eq. (3.35c).

To test the stationarity properties of the spot and futures price series in our sample we employ the third model as the series are not particularly large nor do they exhibit an obvious trend. In our analysis, we test the null hypothesis of a unit root (H_0: α=1) by examination of the Phillips-Perron Z(α) and Z(t_α) test statistics. The statistic Z(t_α) is expressed as:

$$\bar{t}_\alpha = t_\alpha \left(\frac{\gamma_0}{f_0} \right)^{1/2} - \frac{T(f_0 - \gamma_0)(se(\hat{\alpha}))}{2 f_0^{1/2} s} \qquad (3.36)$$

Where, $\hat{\alpha}$ is the estimate of α while t_α is the t-ratio of α, se($\hat{\alpha}$) is the coefficient standard error, and s is the standard error of the test regression. γ_0 is an estimate of the error variance (computed as [T-K]s^2/t, where k is the amount of regressors). f_0 is an estimator of the residual spectrum at a frequency equal to zero.

The error variance is expressed as:

$$y_t = \alpha y_{t-1} + x_t'\delta + \varepsilon_t \tag{3.37}$$

where x is an exogenous regressor, α and δ are estimated parameters and ε_t is the white noise error term.

3.5.3. Panel Unit Root Tests

The stochastic properties of the basis are also tested in a panel setting. Panel unit root tests can be compared to univariate tests. Unit root test are grouped according to restrictions on the AR process across cross-sections of the sample data (Beirne et al., 2007) The panel unit root test as specified by Im et al. (1997) is formulated as:

$$\Delta\tilde{x}_{\text{it}} = \tilde{\mu}_{\text{i0}} + \tilde{\psi}_i \tilde{x}_{\text{it-1}} + \sum_{j=1}^{p-1} \psi_{ij} \tilde{x}_{it-j} + \xi_{\text{it}}. \tag{3.38}$$

where $\bar{x}_{it} = x_{it} - \bar{x}_t,\ \tilde{\mu}_i = \mu_i - \bar{\mu},\ \tilde{\psi}_i = \psi_i - \bar{\psi},\ \tilde{\psi}_{ij} = \psi_{ij} - \bar{\psi}_i, \tilde{\varepsilon}_{it} = \varepsilon_{it} + \theta_t - \bar{\varepsilon}_t,$ θ_t is the time-specific common fixed effect and $\xi_{\text{it}} = \tilde{\varepsilon}_{it} + 1/N \sum_{j=1}^{N} (\psi_i - \psi_j) x_{jt-1}$. Once Eq. (3.38) has been estimated for each cross section of data the t-bar or average Dickey-Fuller test developed by Im et al. (1997) is employed to test the null that each coefficient is non-stationary

$$\bar{t}_{NT} = \frac{1}{N} \sum_{j=1}^{N} t_{iT} \tag{3.39}$$

Where the null assumes the form:

H_0 : ψ_i=0 for i=1…N

H_A : ψ_i <0 for i=1…N

The t-bar statistic $\bar{t}_{NT}$ is then compared with a critical value simulated by Im et al. (1997)

Luintel (2001) applies the test, as formulated by Im et al. (1997), to a panel of 20 OECD real exchange rates against the US dollar and finds that the t-bar test is appropriate to implement for testing the stationarity properties of data exhibiting cross-sectional dependence and dynamic heterogeneity. Luintel (2001) finds that t-bar statistic, $\bar{t}_{11,100} = -2.128$, for the sub-panel of 11 European community members when compared with a 5% critical value of -1.97 rejects the null so that the joint stationarity test is accepted. Alternatively, Hadri (2000) proposes a Lagrange-Multiplier (LM) panel unit root test based on the null of stationarity. Hadri (2000) assumes that the data follows a deterministic path:

$$x_{it} = r_{it} + \varepsilon_{it} \tag{3.40}$$

where *t=1…T* time, *i=1…N* variables. Eq. (3.40) suggests that the series collapses to a random walk and stationary white noise error term:

$$r_{it}=r_{it-1}+u_{it} \tag{3.41}$$

r_{i0} is unknown, u_{it} are independent and identically distributed random variables and $\upsilon_u^2 \geq 0$ where υ_u^2 is the variance term.

The null hypothesis is:

H_0: μ=0 for i=1…N

H_A: μ>0 for i=1…N

where $\mu = \upsilon_u^2 / \upsilon_\varepsilon^2$ and under the assumption that $\sigma_u^2 = 0$ where σ_u^2 is the variance term. Each panel equation can be represented as:

$\mathbf{x_i} = \mathbf{Z_i}\beta_i + \mathbf{e_i}$ (3.42)

where, $\mathbf{x'_i} = [\mathbf{x_{i1}} \ldots \mathbf{x_{iT}}]$, $\mathbf{e'_i} = [\mathbf{e_{i1}} \ldots \mathbf{e_{iT}}]$ and $\mathbf{Z_i}$ is a *Tx1* unit vector. The LM test statistic as formulated by Hadri (2000) to test the null is:

$$LM = \frac{1}{n}\sum_{t=1}^{n} \frac{\frac{1}{T^2}\sum_{t=1}^{T} S_{it}^2}{\sigma_i^{*2}} \quad (3.43)$$

where σ_i^{*2} is the variance estimated from individual samples and the partial sum of the residuals is $S_{it} = \sum_{j=1}^{t} e_{ij^*}$ (3.44)

Expressions of the correction applied to the variance and autocovariance as well as the estimator for the residual spectrum at zero frequency and the Bartlett (BT) and Tukey-Hanning (TH) kernel based estimators, for comparison, as suggested by Hadri (2000) are illustrated in Appendix A3.6. Test statistics based on different kernels, adopted from Beirne et al. (2007), are in addition presented in Appendix A3.6. (see Table A3.16.).

3.6. Empirical Results

3.6.1. Unit Root Tests and Tests of Cointegration

Results for the Phillips-Perron (1988) unit root tests are presented in Tables 3.1. to 3.5. The test statistics confirm that the spot and futures prices, for all five exchange rates, are level non-stationary and that the null of a unit root is rejected for the differenced variables. Augmented Dickey-Fuller (1979) tests are in addition conducted and confirm that all spot rates and futures prices are first difference stationary and therefore integrated of order one I(1), see Appendix A3.2. (Tables A3.1. - A3.5.) for the ADF test results. Since our series are confirmed to be level nonstationary and first difference stationary and thus I(1), we continue by testing for cointegration given the arguments presented earlier.

Table 3.1. Phillips-Perron Unit Root Test Results with a Quadratic Spectral Kernel and Andrews Bandwidth for the USD/GBP Exchange Rate				
$y_t = \alpha y_{t-1} + u_t$ $u_t \sim N(0, \sigma_x^2)$				
Variable	Level		First differences	
	Constant	Constant and trend	Constant	Constant and trend
1 month prior to expiry				
s_t	-2.54 (2.34)	-2.45 (2.32)	-8.26 (0.94)**	-8.25 (0.93)**
f_t	-2.53 (2.38)	-2.43 (2.36)	-8.22 (0.94)**	-8.21 (0.93)**
2 months prior to expiry				
s_t	-2.56 (2.61)	-2.48 (2.59)	-8.00 (1.02)**	-7.98 (1.02)**
f_t	-2.56 (2.5)	-2.48 (2.48)	-8.10 (0.90)**	-8.09 (0.90)**
3 months prior to expiry				
s_t	-2.41 (2.06)	-2.35 (2.05)	-8.64 (0.89)**	-8.61 (0.88)**
f_t	-2.38 (2.04)	-2.32 (2.02)	-8.64 (0.88)**	-8.62 (0.87)**
Notes: the 1% and 5% critical values for the Phillips Perron tests are respectively -3.50 and -2.89 (without trend) and -4.06 and -3.46 (with the trend); ** and * indicate statistical significance at the 1% and 5% levels, respectively. The Andrews bandwidth is presented in brackets ().				

Table 3.2. Phillips Perron Unit Root Test Results with a Quadratic Spectral Kernel aAndrews Bandwidth for the USD/AUD Exchange Rate				
$y_t = \alpha y_{t-1} + u_t$ $u_t \sim N(0, \sigma_x^2)$				
Variable	Level		First differences	
	Constant	Constant and trend	Constant	Constant and trend
1 month prior to expiry				
s_t	-1.68 (1.89)	-2.08 (2.01)	-8.46 (0.49)**	-8.41 (0.49)**
f_t	-1.65 (1.85)	-2.05 (1.97)	-8.50 (0.55)**	-8.45 (0.55)**
2 months prior to expiry				
s_t	-1.58 (1.8)	-1.98 (1.92)	-8.54 (2.81) **	-8.49 (2.91) **
f_t	-1.59 (1.78)	-1.99 (1.9)	-8.56 (0.39) **	-8.51 (0.39) **
3 months prior to expiry				
s_t	-1.55 (2.48)	-2.05 (2.59)	-7.86 (0.57)**	-7.81 (0.56)**
f_t	-1.54 (2.4)	-2.00 (2.5)	-7.95 (0.45)**	-7.90 (0.43)**
Notes: the 1% and 5% critical values for the Phillips Perron tests are respectively -3.50 and -2.89 (without trend) and -4.06 and -3.46 (with the trend); ** and * indicate statistical significance at the 1% and 5% levels, respectively. The Andrews bandwidth is presented in brackets ().				

Table 3.3. Phillips-Perron Unit Root Test Results with a Quadratic Spectral Kernel and Andrews Bandwidth for the USD/CAD Exchange Rate				
$y_t = \alpha y_{t-1} + u_t$ $u_t \sim N(0, \sigma_x^2)$				
Variable	Level		First differences	
	Constant	Constant and trend	Constant	Constant and trend
1 month prior to expiry				
s_t	-1.31 (2.09)	-1.08 (2.12)	-7.97 (0.62)**	-7.97 (0.64)**
f_t	-1.33 (2.07)	-1.11 (2.11)	-7.99 (0.63)**	-7.99 (0.65)**
2 months prior to expiry				
s_t	-1.42 (2.55)	-1.40 (2.61)	-7.54 (0.89) **	-7.52 (0.91) **
f_t	-1.42 (2.68)	-1.38 (2.73)	-7.42 (0.98) **	-7.40 (0.99) **
3 months prior to expiry				
s_t	-1.24 (2.42)	-0.96 (2.43)	-7.39 (0.83)**	-7.43 (0.85)**
f_t	-1.24 (2.45)	-0.96 (2.45)	-7.39 (0.87)**	-7.43 (0.893)**
Notes: the 1% and 5% critical values for the Phillips Perron tests are respectively -3.51 and -2.89 (without trend) and -4.07 and -3.46 (with the trend); ** and * indicate statistical significance at the 1% and 5% levels, respectively. The Andrews bandwidth is presented in brackets ().				

Table 3.4. Phillips Perron Unit Root Test Results with a Quadratic Spectral Kernel and Andrews Bandwidth for the USD/BRL Exchange Rate				
$y_t = \alpha y_{t-1} + u_t$ $u_t \sim N(0, \sigma_x^2)$				
Variable	Level		First differences	
	Constant	Constant and trend	Constant	Constant and trend
1 month prior to expiry				
s_t	-1.89 (1.91)	-1.90 (1.91)	-13.67 (0.47)**	-13.64 (0.47)**
f_t	-1.85 (1.81)	-1.86 (1.82)	-13.76 (0.25)**	-13.73 (0.25)**
2 weeks prior to expiry				
s_t	-1.82 (3.74)	-1.83 (3.74)	-11.31 (1.1)**	-11.28 (1.1)**
f_t	-1.80 (3.99)	-1.82 (3.99)	-11.04 (1.29)**	-11.02 (1.29)**
1 week prior to expiry				
s_t	-1.96 (3.46)	-1.96 (3.46)	-11.65 (0.99)**	-11.62 (0.99)**
f_t	-1.95 (3.4)	-1.95 (3.41)	-11.70 (0.98)**	-11.68 (0.98)**
Notes: the 1% and 5% critical values for the Phillips Perron tests are respectively -3.46 and -2.87 (without trend) and -4.00 and -3.43 (with the trend); ** and * indicate statistical significance at the 1% and 5% levels, respectively. The Andrews bandwidth is presented in brackets ().				

Table 3.5. Phillips Perron Unit Root Test Results with a Quadratic Spectral Kernel and Andrews Bandwidth for the USD/ZAR Exchange Rate				
$y_t = \alpha y_{t-1} + u_t$ $u_t \sim N(0, \sigma_x^2)$				
Variable	Level		First differences	
	Constant	Constant and trend	Constant	Constant and trend
1 Month Prior To Expiry				
s_t	-1.02 (1.36)	-1.53 (1.45)	-14.09 (0.19)**	-14.06 (0.18)**
f_t	-1.04 (1.32)	-1.55 (1.42)	-14.14 (0.19)**	-14.11 (0.19)**
2 Weeks Prior To Expiry				
s_t	-1.13 (1.99)	-1.69 (1.93)	-13.69 (0.57)**	-13.66 (0.57)**
f_t	-1.14 (1.94)	-1.69 (2.01)	-13.58 (0.65)**	-13.56 (0.65)**
1 Week Prior To Expiry				
s_t	-1.02 (0.13)	-1.58 (1.46)	-14.20 (0.44)**	-14.17 (0.43)**
f_t	-1.02 (1.18)	-1.59 (1.3)	-14.34 (0.36)**	-14.31 (0.35)**
Notes: The 1% And 5% Critical Values For The Phillips Perron Tests Are Respectively -3.46 And -2.87 (Without Trend) And -4.00 And -3.43 (With The Trend); ** And * Indicate Statistical Significance At The 1% And 5% Levels, Respectively. The Andrews Bandwidth Is Presented In Brackets ().				

Cointegration is tested by application of the Johansen (1995) methodology discussed in Chapter 2, section 2.3.1. The VAR in error correction form is expressed as follows:

$$\Delta y_t = \mu + \Gamma_1 \Delta y_{t-1} + ... + \Gamma_{k-1} \Delta y_{t-k+1} + \Pi y_{t-k} + \mu_t \quad (3.45)$$

We test the rank of the matrix **Π** which will demonstrate a reduced rank in the instance of no cointegration. Since tests of cointegration are sensitive to the choice of lag length (Hall, 1991) we estimate a VAR (5) to examine the significance of each additional lag. The likelihood ratio tests presented in Appendix A3.2. (Tables A3.6. – A3.10.) confirm that aside from the USD/BRL futures price two weeks prior to expiry, the USD/CAD futures price two months prior to expiry and the AUD/USD futures price two months prior, the appropriate order of the VAR for all the other contracts is one.

The USD/AUD exchange rate rejects restrictions imposed in moving from VAR(2) to VAR(1) two months from expiration, the USD/CAD exchange rate rejects restrictions imposed in moving from VAR(1) to VAR(2) and from VAR(4) to VAR(5) two months from expiration while the USD/BRL rejects restrictions imposed in moving from VAR(1) to VAR(2) two weeks from expiration.

Tests for the number of cointegrating vectors are presented in Tables 3.6.-3.7., based on a VAR with a constant. The deterministic element employed in this instance is directed by the Pantula Principle proposed by Johansen (1992a) and Pantula (1989). The Pantula Principle (Johansen, 1992a and Pantula, 1989) suggests that we begin with the most restrictive model, where r=0 and there is neither an intercept nor a trend in the VAR, and move to the least restrictive model, where r=n-1 and there is both an intercept and a linear trend in the VAR. The trace-test statistics with critical values are compared until for the first time the null hypothesis of no cointegration cannot be rejected. Our tests show that the null hypothesis of zero cointegrating vectors is rejected whilst the null of one cointegrating vector cannot be rejected at the 1% significance level for all VAR models. Since tests of cointegration present evidence of a single cointegrating vector which is a stationary linear combination of two stochastic nonstationary variables (as per the unit root tests in Tables 3.1-3.5) we expect the basis to be path independent.

This finding is significant as it suggests that the markets for spot and futures are efficient and that both price series cointegrate with the cointegrating vector that is the basis. Spot and futures prices under these circumstances will be identical and conditional upon the basis or degree of mispricing resulting from arbitrage activity.

Table 3.6. Tests for the Number of Cointegrating Vectors for the USD/GBP Exchange Rate
$(H_0^a : r = 0, H_1^a : r = 1; H_0^b : r \le 1, H_1^b : r = 2)$

	1 month prior to expiry		2 months prior to expiry		3 months prior to expiry	
	H_0^a	H_0^b	H_0^a	H_0^b	H_0^b	H_0^b
λ_{max}	0.49	0.05	0.43	0.05	0.25	0.06
λ_{trace}	61.25[0.00]**	4.74[0.02]	51.82[0.00]**	4.63[0.03]	29.90[0.00]**	5.95[0.01]
Critical Values (Johansen and Juselius, 1990 Table A.2.)						
			90%		95%	
λ_{max}	r=0		2.81		3.96	
	r=1		12.09		14.03	
λ_{trace}	r=0		2.81		3.96	
	r=1		12.09		15.19	

Table 3.7. Tests for the Number of Cointegrating Vectors for the USD/AUD Exchange Rate
$(H_0^a : r = 0, H_1^a : r = 1; H_0^b : r \le 1, H_1^b : r = 2)$

	1 month prior to expiry		2 months prior to expiry		3 months prior to expiry	
	H_0^a	H_0^b	H_0^a	H_0^b	H_0^b	H_0^b
λ_{max}	0.40	0.02	0.44	0.02	0.31	0.02
λ_{trace}	45.17[0.00]**	2.28[0.13]	51.47[0.00]**	2.20[0.13]	33.98[0.00]**	2.31 [0.12]
Critical Values (Johansen and Juselius, 1990 Table A.2.)						
			90%		95%	
λ_{max}	r=0		2.81		3.96	
	r=1		12.09		14.03	
λ_{trace}	r=0		2.81		3.96	
	r=1		12.09		15.19	

Table 3.8. Tests for the Number of Cointegrating Vectors for the USD/CAD Exchange Rate
$(H_0^a : r = 0, H_1^a : r = 1; H_0^b : r \le 1, H_1^b : r = 2)$

	1 month prior to expiry		2 months prior to expiry		3 months prior to expiry	
	H_0^a	H_0^b	H_0^a	H_0^b	H_0^b	H_0^b
λ_{max}	0.45	0.01	0.27	0.01	0.29	0.01
λ_{trace}	50.03[0.00]**	1.45[0.22]	27.13[0.00]**	1.45[0.22]	1.19[0.27]	28.91[0.00] **
Critical Values (Johansen and Juselius, 1990 Table A.2.)						
			90%		95%	
λ_{max}	r=0		2.81		3.96	
	r=1		12.09		14.03	
λ_{trace}	r=0		2.81		3.96	
	r=1		12.09		15.19	

Table 3.9. Tests for the Number of Cointegrating Vectors for the USD/BRL Exchange Rate $(H_0^a : r=0, H_1^a : r=1; H_0^b : r \le 1, H_1^b : r=2)$						
	1 month prior to expiry		2 weeks prior to expiry		1 week prior to expiry	
	H_0^a	H_0^b	H_0^a	H_0^b	H_0^b	H_0^b
λ_{max}	0.35	0.01	0.46	0.01	0.47	0.01
λ_{trace}	94.97[0.00]**	3.58[0.05]	133.67[0.00]**	2.69[0.10]	141.01[0.00]**	3.41[0.06]
Critical Values (Johansen and Juselius, 1990 Table A.2.)						
			90%		95%	
λ_{max}	r=0		2.81		3.96	
	r=1		12.09		14.03	
λ_{trace}	r=0		2.81		3.96	
	r=1		12.09		15.19	

Table 3.10. Tests for the Number of Cointegrating Vectors for the USD/ZAR Exchange Rate $(H_0^a : r=0, H_1^a : r=1; H_0^b : r \le 1, H_1^b : r=2)$						
	1 month prior to expiry		2 weeks prior to expiry		1 week prior to expiry	
	H_0^a	H_0^b	H_0^a	H_0^b	H_0^b	H_0^b
λ_{max}	0.42	0.00	0.51	0.00	0.44	0.00
λ_{trace}	117.02[0.0]**	0.94[0.33]	155.68[0.0]**	1.10[0.29]	124.13[0.0]**	1.01[0.31]
Critical Values (Johansen and Juselius, 1990 Table A.2.)						
			90%		95%	
λ_{max}	r=0		2.81		3.96	
	r=1		12.09		14.03	
λ_{trace}	r=0		2.81		3.96	
	r=1		12.09		15.19	

3.6.2. The Stochastic Properties of the Basis

As the prerequisite for stationarity have been met both the spot and futures market can be described to function efficiently. We next test the stochastic properties of the basis and expect that it will be I(0). Phillips-Perron (1988) unit root test results for the basis are presented in Tables 3.11. to 3.15. Augmented Dickey-Fuller (1979) tests are in addition conducted for the basis and confirm that it is I(0), see Appendix A3.2. (Tables A3.11. - A3.15.) for the ADF test results.

Table 3.11. Phillips-Perron unit root test results for the basis for the USD/GBP exchange rate (with a Quadratic Spectral Kernel and Andrews Bandwidth)		
$y_t = \alpha y_{t-1} + u_t$ $u_t \sim N(0, \sigma_x^2)$		
Variable	Level	
	Constant	Constant and trend
1 months prior to expiry	-10.32 (0.10)**	-10.24 (1.02)**
2 months prior to expiry	-8.28 (0.79)**	-8.3 (0.77)**
3 months prior to expiry	-4.82 (0.93)**	-4.92 (0.82)**
Note: the 1% and 5% critical values for the Phillips Perron tests are respectively -3.51 and -2.9 (without trend) and -4.07 and -3.46 (with the trend); ** and * indicate statistical significance at the 1% and 5% levels, respectively.		

Table 3.12. Phillips-Perron unit root test results for the basis for the USD/AUD exchange rate (with a Quadratic Spectral Kernel and Andrews Bandwidth)		
$y_t = \alpha y_{t-1} + u_t$ $u_t \sim N(0, \sigma_x^2)$		
Variable	Level	
	Constant	Constant and trend
1 month prior to expiry	-7.41 (0.85)**	-7.63 (0.82)**
2 months prior to expiry	-7.75 (0.8) **	-8.54 (0.67) **
3 month prior to expiry	-4.42 (1.96)**	-5.51 (1.74)**
Note: the 1% and 5% critical values for the Phillips Perron tests are respectively -3.51 and -2.9 (without trend) and -4.07 and -3.46 (with the trend); ** and * indicate statistical significance at the 1% and 5% levels, respectively.		

Table 3.13. Phillips-Perron unit root test results for the basis for the USD/CAD exchange rate (with a Quadratic Spectral Kernel and Andrews Bandwidth)		
$y_t = \alpha y_{t-1} + u_t$ $u_t \sim N(0, \sigma_x^2)$		
Variable	Level	
	Constant	Constant and trend
1 month prior to expiry	-7.84 (0.22)**	-8.05 (0.35)**
2 month prior to expiry	-5.39 (1.78) **	-5.59 (1.75) **
3 month prior to expiry	-5.67 (1.66)**	-5.83 (1.64)**
Note: the 1% and 5% critical values for the Phillips Perron tests are respectively -3.51 and -2.9 (without trend) and -4.07 and -3.47 (with the trend); ** and * indicate statistical significance at the 1% and 5% levels, respectively.		

Table 3.14. Phillips-Perron unit root test results for the basis for the USD/BRL exchange rate (with a Quadratic Spectral Kernel and Andrews Bandwidth)		
$y_t = \alpha y_{t-1} + u_t$ $u_t \sim N(0, \sigma_x^2)$		
Variable	Level	
	Constant	Constant and trend
1 month prior to expiry	-10.40 (1.43)**	-10.55 (1.38)**
2 weeks prior to expiry	-12.88 (7.69)**	-13.06 (6.73)**
1 week prior to expiry	-13.82 (0.395)**	-13.82 (0.41)**
Note: the 1% and 5% critical values for the Phillips Perron tests are respectively -3.46 and -2.87 (without trend) and -4.00 and -3.43 (with the trend); ** and * indicate statistical significance at the 1% and 5% levels, respectively.		

Table 3.15. Phillips-Perron unit root test results for the basis for the USD/ZAR exchange rate (with a Quadratic Spectral Kernel and Andrews Bandwidth)		
$y_t = \alpha y_{t-1} + u_t$ $u_t \sim N(0, \sigma_x^2)$		
Variable	Level	
	Constant	Constant and trend
1 month prior to expiry	-12.28 (0.69)**	-12.36 (0.74)**
2 weeks prior to expiry	-15.12 (0.19)**	-15.19 (0.20)**
1 week prior to expiry	-12.91 (1.01)**	-12.88 (1.01) **
Note: the 1% and 5% critical values for the Phillips Perron tests are respectively -3.46 and -2.88 (without trend) and -4.00 and -3.43 (with the trend); ** and * indicate statistical significance at the 1% and 5% levels, respectively.		

Cointegrating vectors are estimated using the Johansen (1995) methodology and are presented below in Tables 3.16.- 3.20. These results suggest that prices are identical in the spot and futures market and contingent on the basis. The cointegration tests suggest that s_t-f_t is an I(0) series. The reduced form system is identified as a VAR(1) from likelihood ratio tests for VAR length. This identification with confirmation that the VAR(1) is the basis confirms that the re-parameterized VAR(1) in error correction form is the most appropriate model.

Table 3.16. The Estimated Cointegrating Vector for the Basis of the USD/GBP Exchange Rate			
	1 month prior to expiry	2 months prior to expiry	3 months prior to expiry
coefficient	0.99	1.00	1.00
Standard error	(0.00)	(0.00)	(0.00)
t-stat	342.44**	230.88**	145.85**
Notes: ** and * indicate significance at the 1% and 5% levels, respectively based on standard inference.			

Table 3.17. The Estimated Cointegrating Vector for the Basis of the USD/AUD Exchange Rate			
	1 month prior to expiry	2 months prior to expiry	3 months prior to expiry
coefficient	0.99	0.99	0.99
Standard error	(0.00)	(0.00)	(0.00)
t-stat	376.16**	414.58**	284.85**
Notes: ** and * indicate significance at the 1% and 5% levels, respectively based on standard inference.			

Table 3.18. The Estimated Cointegrating Vectors for the Basis of the USD/CAD Exchange Rate			
	1 month prior to expiry	2 months prior to expiry	3 months prior to expiry
coefficient	1.00	1.00	1.00
Standard error	(0.00)	(0.00)	(0.00)
t-stat	500**	333.33**	333.33**
Notes: ** and * indicate significance at the 1% and 5% levels, respectively based on standard inference.			

Table 3.19. The Estimated Cointegrating Vector for the Basis of the USD/BRL Exchange Rate			
	1 month prior to expiry	2 weeks prior to expiry	1 week prior to expiry
coefficient	0.99	1.00	0.99
Standard error	(0.00)	(0.00)	(0.00)
t-stat	330**	1000**	434.23**
Notes: ** and * indicate significance at the 1% and 5% levels, respectively based on standard inference.			

Table 3.20. The Estimated Cointegrating Vector for the Basis of the USD/ZAR Exchange Rate			
	1 month prior to expiry	2 weeks prior to expiry	1 week prior to expiry
coefficient	0.99	0.99	1.00
Standard error	(0.00)	(0.00)	(0.00)
t-stat	623.61**	818.85**	588.23**
Notes: ** and * indicate significance at the 1% and 5% levels, respectively based on standard inference.			

The cointegrating vectors (i.e. the basis) plotted at 1-month, 2-month and 3-month intervals prior to expiry for the USD vis-à-vis the GBP, AUD and CAD, correspondingly are shown in Appendix A3.4. (Figures A3.14.-3.22.). We in addition present cointegrating vectors (i.e. the basis) plotted at 1-week, 2-weeks and 1-month intervals prior to expiry for the USD vis-à-vis the BRL and ZAR in Appendix A3.4. (figures A3.23.-3.28.). We recall that the price series are constructed as a sequence rather than a series. The price series (spot and futures rates) quoted for each nearest maturity contract is quarterly for the AUD, CAD and GBP vis-à-vis the USD and monthly for the BRL and ZAR vis-à-vis the USD.

3.6.3. Panel Unit Root Test Results

Results of panel unit root tests for the basis are presented in Tables 3.22. – 3.26. We observe that the null of non-stationarity is rejected by test statistics at the 1% level for the basis relating to all five exchange rates. The Levin et al. (2002) panel test that assumes independent identically distributed innovations is in addition applied to provide a benchmark and comparability to results from the Im et al. (1997) and Hadri (2000) tests. Results of the Hadri (2000) test cannot reject the null of stationarity for the basis so confirm results of the Levin et al. (2002) and Im et al. (1997) panel unit root tests.

Results of the individual unit root tests as well as the panel unit root tests confirm our earlier intuition that the basis is indeed stochastic.

Table 3.22. Panel Unit Root Tests for the USD/GBP Basis				
Method	Level			
	Constant		Constant and trend	
	Statistic	Probability	Statistic	Probability
Levin, Lin & Chu t*	-9.38	0.00	-9.64	0.00
Im, Pesaran and Shin W-stat	-12.67	0.00	-12.58	0.00
Hadri test				
Hadri Z-stat	-0.06	0.52	0.93	0.17
Heteroscedastic Consistent Z-stat	-0.18	0.57	0.99	0.16

Table 3.23. Panel Unit Root Tests for the USD/AUD Basis				
Method	Level			
	Constant	Constant and trend		
	Statistic	Probability	Statistic	Probability
Levin, Lin & Chu t*	-6.04	0.00	-7.01	0.00
Im, Pesaran and Shin W-stat	-8.90	0.00	-9.60	0.00
Hadri test				
Hadri Z-stat	1.50	0.06	0.03	0.48
Heteroscedastic Consistent Z-stat	1.23	0.10	-0.11	0.54

Table 3.24. Panel Unit Root Tests for the USD/CAD Basis				
Method	Level			
	Constant		Constant and trend	
	Statistic	Probability	Statistic	Probability
Levin, Lin & Chu t*	-3.69	0.00	-3.63	0.00
Im, Pesaran and Shin W-stat	-6.78	0.00	-6.52	0.00
Hadri test				
Hadri Z-stat	-0.18	0.57	-0.08	0.53
Heteroscedastic Consistent Z-stat	0.19	0.43	0.39	0.35

Table 3.25. Panel Unit Root Tests for the USD/BRL Basis				
Method	Level			
	Constant		Constant and trend	
	Statistic	Probability	Statistic	Probability
Levin, Lin & Chu t*	-9.46	0.00	-15.24	0.00
Im, Pesaran and Shin W-stat	-14.11	0.00	-17.70	0.00
Hadri test				
Hadri Z-stat	1.29	0.09	1.07	0.14
Heteroscedastic Consistent Z-stat	1.15	0.12	1.00	0.15

Table 3.26. Panel unit root tests for the USD/ZAR basis				
Method	Level			
	Constant		Constant and trend	
	Statistic	Probability	Statistic	Probability
Levin, Lin & Chu t*	-17.21	0.00	-20.26	0.00
Im, Pesaran and Shin W-stat	-20.07	0.00	-21.22	0.00
Hadri test				
Hadri Z-stat	-0.10	0.54	-0.22	0.58
Heteroscedastic Consistent Z-stat	-0.03	0.51	0.07	0.47

3.7. Conclusions

In this chapter, we have analyzed the pricing relationship as well as the efficiency of the spot and futures market for foreign exchange using data for the GBP, AUD, CAD, BRL and ZAR vis-à-vis the USD. To test market efficiency, we examined the unit root properties of the spot and futures prices at several intervals to expiry. The null of a unit root could not be rejected for the spot and futures prices for any of the exchange rates in our sample. We further our analysis by application of the Johansen (1995) procedure for cointegration and find that the spot and futures prices do cointegrate and that the cointegrating vector is the basis. In addition, the finding of cointegration accommodates the first condition for effectively functioning markets.

We identify the reduced form system as a VAR(1) from Likelihood Ratio tests for VAR length. This identification with confirmation that the cointegrating vector is the basis infers that the VAR(1) reparametrized in error correction form is the most apt model specification. Subsequently, the second condition for identification is in addition accommodated. We therefore deduce that both markets are efficient since these conditions have been met, no matter the interval to maturity.

We lastly scrutinize the behaviour of the basis. Results of the individual unit root tests for the basis cannot reject non-stationarity so we confirm that the linear combination of the spot and futures prices, that are individually integrated of order one I(1) are a stationary stochastic process that is the basis. We next test the unit root properties of the basis in a panel. The Im et al. (1997) test which pools as well as demeans data, yields results that are consistent with those of Luintel (2001). The Hadri (2002) test which is powerful and can be corrected for serial correlation and heteroscedasticity confirms the results of Im et al. (1997) that the basis is indeed a stationary unit root process.

In this paper, we find conclusive evidence that the spot and futures market for the GBP, CAD, AUD, BRL and ZAR vis-à-vis the USD are efficient. The inference of this finding is that the futures market is fulfilling its function to provide risk free arbitrage opportunities and that the spot market accomplishes its function as an appropriator of means.

We contribute to the extant literature by expanding on the analysis of Antoniou and Garrett (1993) with relation to futures contracts on the FTSE to operating of the spot and futures market for foreign currency. We find evidence by employing the Johansen (1995) methodology of a cointegrating relation for each currency pair. Implying that under cointegration the basis is path independent. We observe that unit root tests confirm that the basis for all currency pairs is stationary. We cannot reject stationarity when applied to pooled data using the panel tests of Levin et al. (2002) and Im et al. (1997) under the null of non-stationarity & Hadri (2000) under the null of stationarity. We believe that there is scope to extend this study to other emerging market currencies in view of more accessible coherent data and a growing market share. We in addition recommend that this study be extended to other financial markets and that the notion that futures market volatility affects the market for spot be examined.

Appendix A3.1.

The Nature of Futures Contracts

Futures contracts are deferred claims on assets and to discern from forwards are standardized contracts traded on an organized exchange. Futures contracts are fixed in size, delivery date, place of delivery and contract commitments. Regulatory institutions record futures' trades and clearing houses insure against default by reconciling gains and losses daily (Hull, 2015). Clearing houses contribute to smooth functioning of the futures market and act to reduce transaction costs. In order to trade on a market for futures a trader must post a margin or deposit with a broker. Trade losses below the margin threshold are reconciled daily so that the margin acts as a deposit against default. Foreign currency futures contribute to establishing a perfect market by hedging market risk so that investors benefit from a range of opportunities.

Since a futures contract is a delayed claim on an asset it must intrinsically hold information about the future. We evaluate the extent to which futures deliver accurate information about spot rates. Standardized contracts, consolidated clearing houses, posting of a margin (possibly an interest-bearing asset) and reduced transaction costs allow for swift transactions and quick dissemination of information to hedgers, arbitrageurs and speculators. Speculative market participants are observed to undermine the spot market and generate volatility in prices thereby transmitting false cues to hedgers and arbitrageurs. A conflicting statement has been made regarding the UK futures market where speculators are reported to buy low and sell high averting large price fluctuations thereby stabilizing underlying spot prices (Antoniou and Foster, 1992; Antoniou and Holmes, 1992).

Futures and Market Completeness

The market for futures allows investors to defer claims on assets so that current investments are made with a view to increased consumption in the future. Since expectations about the future are unsure, investors incur an element of risk in the investment-consumption decision. The return on an investment relies on a future world predicament and therein lies an undetermined element of risk. Investors will seek to eliminate future risk but in a precarious world where a market for all securities is non-existent an investor cannot shape future returns so as to encapsulate every world predicament. Since investor preferences vary and a market that caters to all is non- existent the opportunity to share risk is negligible. Futures contracts help to complete the market since they allow investors to contemplate future world predicaments whilst sharing risk efficiently and are less costly to construct than ‘primitive’ (Ross, 1976a) securities.

The Market for Foreign Exchange Futures

Currency futures contracts were first traded in an open outcry on the trading floor of the Chicago Mercantile Exchange (CME) but in keeping with the times an electronic trading platform was launched, CME GLOBEX, in 1992. Electronic trading is at present the most common medium for trading futures contracts. Market participants can view their buy and sell orders as well as those of other traders so that the dissemination of information is free and instantaneous. Electronic trading platforms have enhanced the extent and speed at which assets internalize information into prices and have thus improved market efficiency despite the persistence of volatility (Balduzzi et al. 2001).

Volatility persistence has been linked to the delivery of information in clusters (Clark, 1973 and Harris, 1987), autocorrelation induced by nonsynchronous market trading (Harris, 1989; Romer, 1993; Hsieh and Kleidon, 1996; Andersen and Bollerslev, 1998), liquidity traders who wish to capitalize on market liquidity following macroeconomic news (Admati and Pfleiderer, 1988; Ederington and Lee, 1993, 1995) and information asymmetry arising from the varying skills of market participants to decipher information (Kim and Verrecchia, 1994).

Foreign currency futures contracts are most commonly traded on the International Currency Exchange (ICE), the Intercontinental Exchange, Euronext LIFFE, the Tokyo Financial Exchange as well as the CME. Futures contracts are typically delivered quarterly on the third Wednesday in March, June, September and December. The price of a futures contract is dictated by supply and demand so that short-selling would reduce prices and vice versa. Commencement of trading is relayed by the relevant exchange and sojourns a couple of days prior to the last delivery date. Currency futures are generally quoted in terms of a foreign currency against the USD. Exchanges enforce price limits on contracts to avoid large price fluctuations and speculative trading. As another measure to avert speculative trading, investors are restricted by the exchange to hold no more than a fixed maximum of contracts (position limit). The delivery date is the contract's expiry and the contract is settled in cash at the closing price on the expiry date. Traders either buy or sell contingent upon their opening position and stand to lose or gain subject to the disparity between opening and closing prices. Currency futures therefore provide investors with an opportunity to create returns that align with future aspirations.

Economic Benefits of Foreign Exchange Futures

Foreign exchange future contracts are applied by hedgers, speculators and arbitrageurs. Currency futures enhance market efficiency and provide a space where traders can accommodate each other's needs. Hedgers take a position (buy or sell) in the futures market to guarantee against negative repercussions of a planned transaction on the spot market. Futures allow hedgers to pass risk on to speculators who stand to gain as risk increases and premiums rise. Speculators supply the futures market with liquidity and stand to profit from volatility and frequent trading as they generally assume short positions. Arbitrageurs are perhaps the most significant players in the futures market as they align prices by assuming simultaneous positions. Arbitrageurs profit from speculative trade and would sell overvalued futures to buy spot thereby shifting stress from the futures market to the spot market subsequently realigning prices. Hedgers look to arbitrageurs for cues so that their current position compensates for future deficit or the reverse.

Appendix A3.2.

Tables A3.1. - A3.5. ADF unit root test results for spot and futures prices

Table A3.1. ADF Unit Root Test Results Based on SBIC for the USD/GBP Exchange Rate				
$\Delta x_t = \pi_0 + \gamma x_{t-1} + \pi_1 \Delta x_{t-1} + u_t$ $u_t \sim N(0, \sigma_x^2)$				
Variable	Level		First differences	
	Constant	Constant and trend	Constant	Constant and trend
1 month prior to expiry				
s_t	-2.36 (0)	-2.26 (0)	-8.26 (0)**	-8.25 (0)**
f_t	-2.34 (0)	-2.23 (0)	-8.21 (0)**	-8.21 (0)**
2 months prior to expiry				
s_t	-2.35 (0)	-2.27 (0)	-7.99 (0)**	-7.97 (0)**
f_t	-2.36 (0)	-2.26 (0)	-8.10 (0)**	-8.09 (0)**
3 months prior to expiry				
s_t	-2.29 (0)	-2.22 (0)	-8.64 (0)**	-8.61 (0)**
f_t	-2.26 (0)	-2.18 (0)	-8.65 (0)**	-8.62 (0)**
Note: the 1% and 5% critical values for the ADF tests are respectively -3.5 and -2.89 (without trend) and -4.06 and -3.46 (with trend); allowing for a maximum of 14 lags; ** and * indicate statistical significance at the 1% and 5% levels, respectively. Lag length is in brackets ().				

Table A3.2. ADF Unit Root Test Results Based on SBIC for the USD/AUD Exchange Rate				
$\Delta x_t = \pi_0 + \gamma x_{t-1} + \pi_1 \Delta x_{t-1} + u_t$ $u_t \sim N(0, \sigma_x^2)$				
Variable	Level		First differences	
	Constant	Constant and trend	Constant	Constant and trend
1 month prior to expiry				
s_t	-1.57 (0)	-1.92 (0)	-8.46 (0) **	-8.41 (0)**
f_t	-1.55 (0)	-1.90 (0)	-8.50 (0) **	-8.45 (0)**
2 months prior to expiry				
s_t	-1.48 (0)	-1.83(0)	-8.54 (0) **	-8.49 (0) **
f_t	-1.49 (0)	-1.84 (0)	-8.56 (0) **	-8.51 (0) **
3 months prior to expiry				
s_t	-1.37 (0)	-1.82 (0)	-7.86 (0) **	-7.81 (0) **
f_t	-1.37 (0)	-1.78 (0)	-7.95 (0) **	-7.90 (0) **
Note: the 1% and 5% critical values for the ADF tests are respectively -3.5 and -2.89 (without trend) and -4.06 and -3.46 (with trend); allowing for a maximum of 14 lags; ** and * indicate statistical significance at the 1% and 5% levels, respectively. Lag length is in brackets ().				

Table A3.3. ADF Unit Root Test Results Based on SBIC for the USD/CAD Exchange Rate

$$\Delta x_t = \pi_0 + \gamma x_{t-1} + \pi_1 \Delta x_{t-1} + u_t$$

$$u_t \sim N(0, \sigma_x^2)$$

Variable	Level		First differences	
	Constant	Constant and trend	Constant	Constant and trend
1 month prior to expiry				
s_t	-1.24 (0)	-0.88 (0)	-7.97 (0)**	-7.97 (0)**
f_t	-1.25 (0)	-0.90 (0)	-7.99 (0)**	-7.99 (0)**
2 months prior to expiry				
s_t	-1.29 (0)	-1.05(0)	-7.55 (0)**	-7.53 (0)**
f_t	-1.28 (0)	-1.00 (0)	-7.41 (0)**	-7.39 (0)**
3 months prior to expiry				
s_t	-1.15 (0)	-0.69 (0)	-7.41 (0)**	-7.44 (0)**
f_t	-1.14 (0)	-0.69 (0)	-7.39 (0)**	-7.43 (0)**

Note: the 1% and 5% critical values for the ADF tests are respectively -3.5 and -2.89 (without trend) and -4.06 and -3.46 (with trend); allowing for a maximum of 14 lags; ** and * indicate statistical significance at the 1% and 5% levels, respectively. Lag length is in brackets ().

Table A3.4. ADF Unit Root Test Results Based on SBIC for the USD/BRL Exchange Rate

$$\Delta x_t = \pi_0 + \gamma x_{t-1} + \pi_1 \Delta x_{t-1} + u_t$$

$$u_t \sim N(0, \sigma_x^2)$$

Variable	Level		First differences	
	Constant	Constant and trend	Constant	Constant and trend
1 month prior to expiry				
s_t	-1.82 (0)	-1.83 (0)	-13.67 (0)**	-13.64 (0)**
f_t	-1.79 (0)	-1.79 (0)	-13.76 (0)**	-13.73 (0)**
2 weeks prior to expiry				
s_t	-1.96 (1)	-1.97 (1)	-11.29 (0)**	-11.26 (0)**
f_t	-1.95 (1)	-1.96 (1)	-10.98 (0)**	-10.96 (0)**
1 week prior to expiry				
s_t	-2.13 (1)	-2.13 (1)	-11.63 (0)**	-11.61 (0)**
f_t	-2.12 (1)	-2.12 (1)	-11.69 (0)**	-11.67 (0)**

Note: the 1% and 5% critical values for the ADF tests are respectively -3.5 and -2.89 (without trend) and -4.06 and -3.46 (with the trend); allowing for a maximum of 14; ** and * indicate statistical significance at the 1% and 5% levels, respectively. Lag length is in brackets ().

Table A3.5. ADF Unit Root Test Results Based on SBIC for the USD/ZAR Exchange Rate				
$\Delta x_t = \pi_0 + \gamma x_{t-1} + \pi_1 \Delta x_{t-1} + u_t$ $u_t \sim N(0, \sigma_x^2)$				
Variable	Level		First differences	
	Constant	Constant and trend	Constant	Constant and trend
1 month prior to expiry				
s_t	-0.99 (0)	-1.50 (0)	-14.09 (0)**	-14.06 (0)**
f_t	-1.01 (0)	-1.52 (0)	-14.14 (0)**	-14.11 (0)**
2 weeks prior to expiry				
s_t	-1.05 (0)	-1.59 (0)	-13.69 (0)**	-13.66 (0)**
f_t	-1.06 (0)	-1.60 (0)	-13.58 (0)**	-13.56 (0)**
1 week prior to expiry				
s_t	-0.98 (0)	-1.53 (0)	-14.20 (0)**	-14.17 (0)**
f_t	-1.00 (0)	-1.55 (0)	-14.34 (0)**	-14.31(0)**
Note: the 1% and 5% critical values for the ADF tests are respectively -3.5 and -2.89 (without trend) and -4.06 and -3.46 (with trend); allowing for a maximum of 14 lags; ** and * indicate statistical significance at the 1% and 5% levels, respectively. Lag length is in brackets ().				

Tables A3.6 – A3.10. Likelihood Ratio Tests for VAR Length (Distributed χ^2 (4))

Table A3.6. Likelihood ratio tests for VAR length (distributed $x^2(4)$) for the USD/GBP exchange rate			
	1 month prior to expiry	2 months prior to expiry	3 months prior to expiry
5 to 4	2.79 [0.02]*	1.25 [0.29]	1.53 [0.19]
4 to 3	0.46 [0.76]	0.69 [0.59]	0.92 [0.45]
3 to 2	2.11 [0.08]	1.06 [0.37]	0.72 [0.57]
2 to 1	2.24 [0.06]	2.65 [0.03]*	0.42 [0.79]
1 to 0	18.31 [0.00]**	19.60 [0.00]**	23.28 [0.00]**

Table A3.7. Likelihood ratio tests for VAR length (distributed $x^2(4)$) for the USD/AUD exchange rate			
	1 month prior to expiry	2 months prior to expiry	3 months prior to expiry
5 to 4	1.43 [0.22]	0.02 [0.99]	0.45 [0.76]
4 to 3	0.22 [0.92]	1.18 [0.32]	0.52 [0.71]
3 to 2	0.90 [0.46]	5.61 [0.00]**	0.32 [0.86]
2 to 1	0.10 [0.98]	0.43 [0.77]	1.58 [0.18]
1 to 0	17.08 [0.00]**	13.46 [0.00]**	18.36 [0.00]**

Table A3.8. Likelihood ratio tests for VAR length (distributed $x^2(4)$) for the USD/CAD exchange rate			
	1 month prior to expiry	2 month prior to expiry	3 months prior to expiry
5 to 4	1.42 [0.23]	4.68 [0.00]**	0.81 [0.52]
4 to 3	0.78 [0.53]	0.93 [0.44]	0.41 [0.79]
3 to 2	0.15 [0.96]	2.98 [0.02]*	0.95 [0.43]
2 to 1	0.41 [0.80]	3.54 [0.00]**	1.94 [0.10]
1 to 0	16.65 [0.00]**	25.48 [0.00]**	19.00 [0.00]**

Table A3.9. Likelihood ratio tests for VAR length (distributed $x^2(4)$) for the USD/BRL exchange rate			
	1 month prior to expiry	2 weeks prior to expiry	1 week prior to expiry
5 to 4	1.93 [0.10]	0.43 [0.78]	0.91 [0.45]
4 to 3	2.18 [0.07]	0.32 [0.86]	1.66 [0.15]
3 to 2	0.34 [0.84]	1.01 [0.39]	0.72 [0.57]
2 to 1	3.23 [0.01]*	3.46 [0.00]**	2.46 [0.04]*
1 to 0	50.16 [0.00]**	64.05 [0.00]**	58.63 [0.00]**

Table A3.10. Likelihood ratio tests for VAR length (distributed $x^2(4)$) for the USD/ZAR exchange rate			
	1 month prior to expiry	2 weeks prior to expiry	1 week prior to expiry
5 to 4	2.06 [0.08]	0.47 [0.75]	0.30 [0.87]
4 to 3	0.07 [0.98]	0.29 [0.88]	0.15 [0.96]
3 to 2	0.31 [0.86]	0.72 [0.57]	0.79 [0.52]
2 to 1	0.30 [0.87]	1.23 [0.29]	2.15 [0.07]
1 to 0	44.30 [0.00]**	46.08 [0.00]**	44.39 [0.00]**

Tables A3.11.- A3.15. ADF unit root tests for the basis

Table A3.11. ADF Unit Root Test Results Based on SBIC for the Basis for the USD/GBP Exchange Rate		
$\Delta x_t = \pi_0 + \gamma x_{t-1} + \pi_1 \Delta x_{t-1} + u_t$ $u_t \sim N(0, \sigma_x^2)$		
Variable	Level	
	Constant	Constant and trend
1 month prior to expiry	-10.32 (0)**	-10.24 (0)**
2 months prior to expiry	-8.28 (0)**	-8.29 (0)**
3 months prior to expiry	-4.81 (0)**	-4.92 (0)**
Note: the 1% and 5% critical values for the ADF tests are respectively -3.51 and -2.9 (without trend) and -4.07 and -3.46 (with trend); allowing for a maximum of four lags; ** and * indicate statistical significance at the 1% and 5% levels, respectively. Lag length is in brackets ().		

Table A3.12. ADF Unit Root Test Results Based on SBIC for the Basis for the USD/AUD Exchange Rate		
$\Delta x_t = \pi_0 + \gamma x_{t-1} + \pi_1 \Delta x_{t-1} + u_t$ $u_t \sim N(0, \sigma_x^2)$		
Variable	Level	
	Constant	Constant and trend
1 month prior to expiry	-7.40 (0)**	-7.62 (0)**
2 months prior to expiry	-7.75 (0) **	-8.54 (0) **
3 months prior to expiry	-2.70 (1)	-3.34 (1)
Note: the 1% and 5% critical values for the ADF tests are respectively -3.51 and -2.9 (without trend) and -4.07 and -3.46 (with the trend); the proper lag length allowing for a maximum of four lags is selected by the Schwarz Bayesian information criterion (SBIC) ** and * indicate statistical significance at the 1% and 5% levels, respectively. Lag length is in brackets ().		

Table A3.13. ADF Unit Root Test Results Based on SBIC for the Basis for the USD/CAD Exchange Rate		
$\Delta x_t = \pi_0 + \gamma x_{t-1} + \pi_1 \Delta x_{t-1} + u_t$ $u_t \sim N(0, \sigma_x^2)$		
Variable	Level	
	Constant	Constant and trend
1 month prior to expiry	-7.84 (0)**	-8.054 (0)**
2 month prior to expiry	-2.30(4)	-2.54 (4)
3 months prior to expiry	-3.68 (1)**	-3.91(1)*
Note: the 1% and 5% critical values for the ADF tests are respectively -3.47 and -2.88 (without trend) and -4.02 and -3.44 (with the trend); the proper lag length allowing for a maximum of four lags is selected by the Schwarz Bayesian information criterion (SBIC) ** and * indicate statistical significance at the 1% and 5% levels, respectively. Lag length is in brackets ().		

Table A3.14. ADF Unit Root Test Results Based on SBIC for the Basis for the USD/BRL Exchange Rate		
$\Delta x_t = \pi_0 + \gamma x_{t-1} + \pi_1 \Delta x_{t-1} + u_t$ $u_t \sim N(0, \sigma_x^2)$		
Variable	Level	
	Constant	Constant and trend
1 month prior to expiry	-7.36 (1)**	-10.64 (0)**
2 weeks prior to expiry	-12.88 (0)**	-13.06 (0)**
1 week prior to expiry	-13.82 (0)	-13.82 (0)*
Note: the 1% and 5% critical values for the ADF tests are respectively -3.46 and -2.88 (without trend) and -4.00 and -3.43 (with the trend); the proper lag length allowing for a maximum of four lags is selected by the Schwarz Bayesian information criterion (SBIC) ** and * indicate statistical significance at the 1% and 5% levels, respectively. Lag length is in brackets ().		

Table A3.15. ADF Unit Root Test Results Based on SBIC for the Basis for the USD/ZAR Exchange Rate		
$\Delta x_t = \pi_0 + \gamma x_{t-1} + \pi_1 \Delta x_{t-1} + u_t$ $u_t \sim N(0, \sigma_x^2)$		
Variable	Level	
	Constant	Constant and trend
1 month prior to expiry	-12.28 (0)**	-12.36 (0)**
2 weeks prior to expiry	-15.12 (0) **	-15.19 (0) **
1 week prior to expiry	-12.91 (0) **	-12.88 (0) **
Note: the 1% and 5% critical values for the ADF tests are respectively -3.46 and -2.88 (without trend) and -4.00 and -3.43 (with the trend); the proper lag length allowing for a maximum of four lags is selected by the Schwarz Bayesian information criterion (SBIC) ** and * indicate statistical significance at the 1% and 5% levels, respectively. Lag length is in brackets ().		

Appendix A3.3.

Figures A3.1. - A3.13. Graphical Representation of the Spot and Futures Prices Shadowing Each Other

Figure A3.1. The USD/GBP Spot Exchange Rate and Futures Prices 3 Months Prior to Expiry

Figure A3.2. The USD/GBP Spot Exchange Rate and Futures Prices 2 Months Prior to Expiry

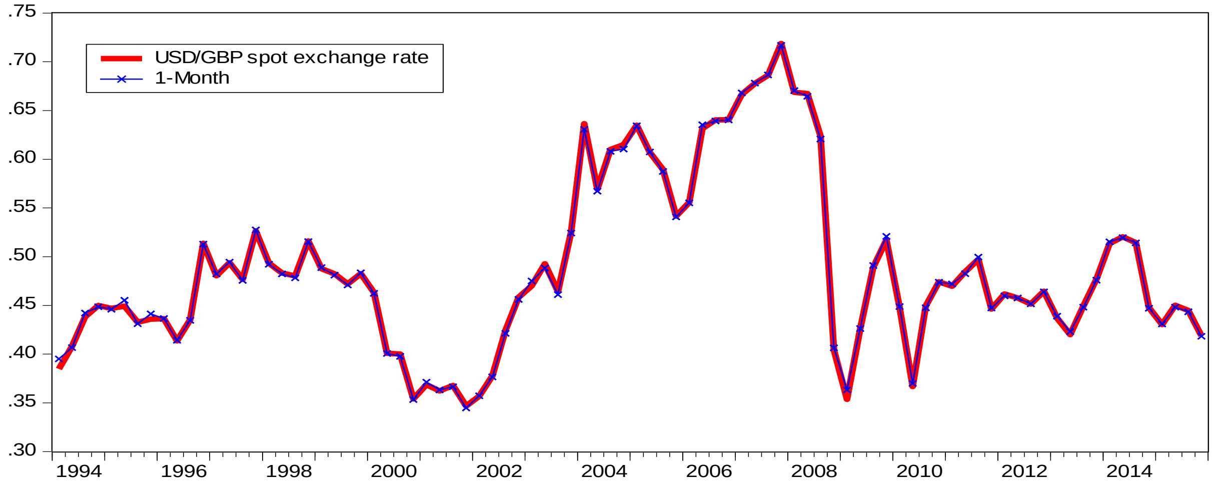

Figure A3.3. The USD/GBP Spot Exchange Rate and Futures Prices 1 Month Prior to Expiry

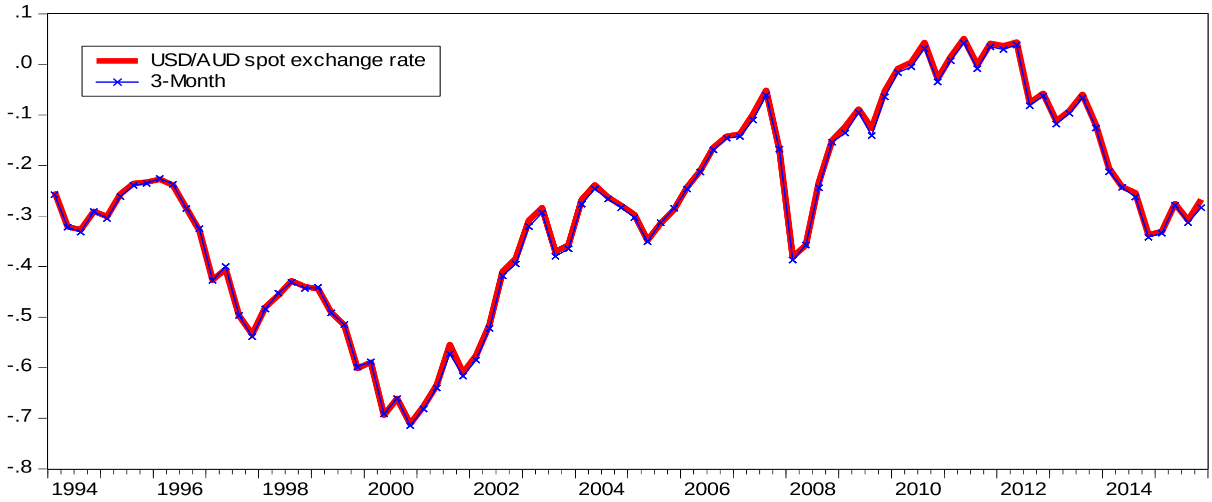

Figure A3.4. The USD/AUD Spot Exchange Rate and Futures Prices 3 Months Prior to Expiry

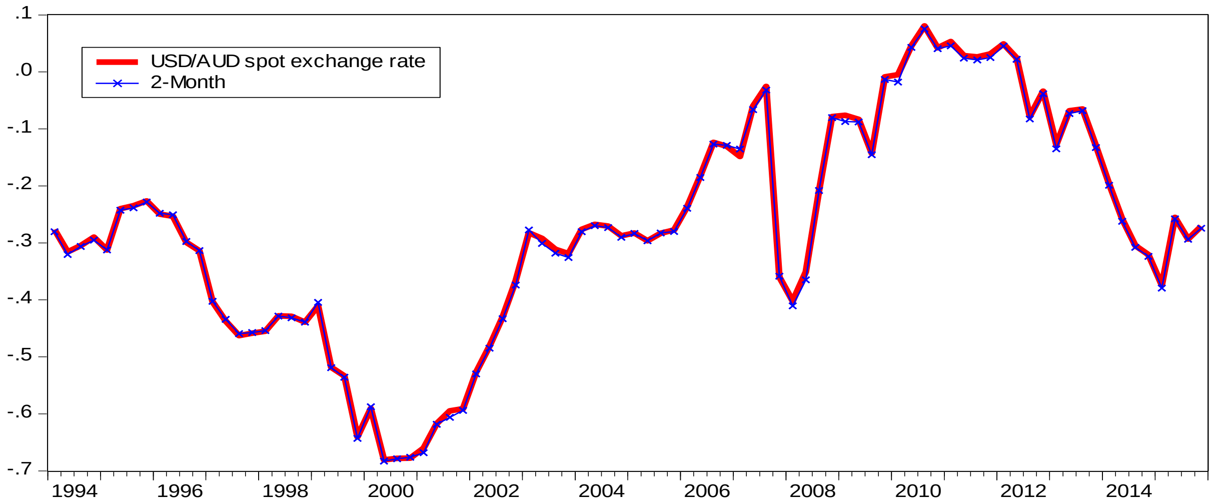

Figure A3.5. The USD/AUD Spot Exchange Rate and Futures Prices 2 Months Prior to Expiry

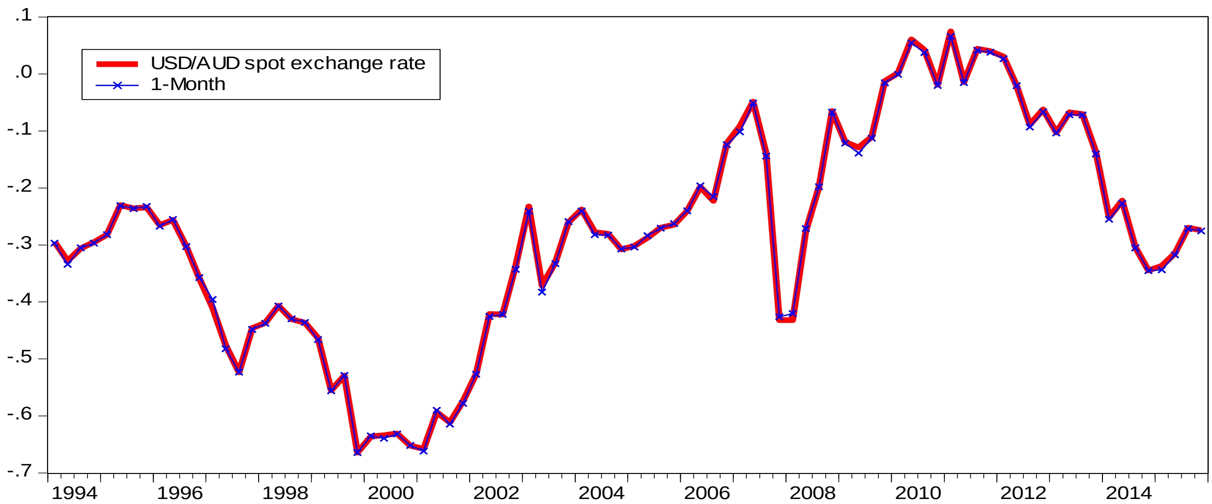

Figure A3.6. The USD/AUD Spot Exchange Rate and Futures Prices 1 Months Prior to Expiry

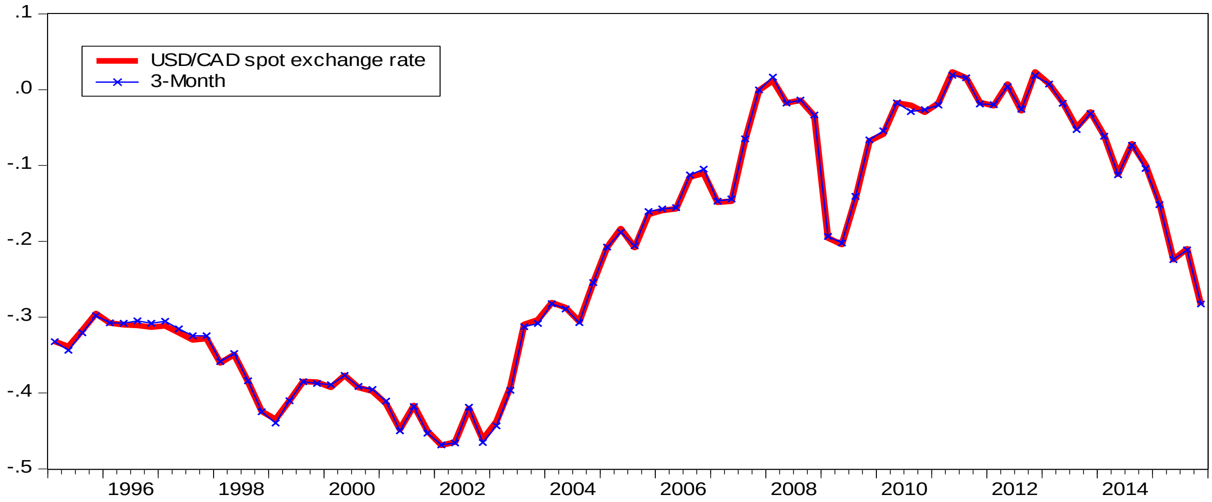

Figure A3.7. The USD/CAD Spot Exchange Rate and Futures Prices 3 Months Prior to Expiry

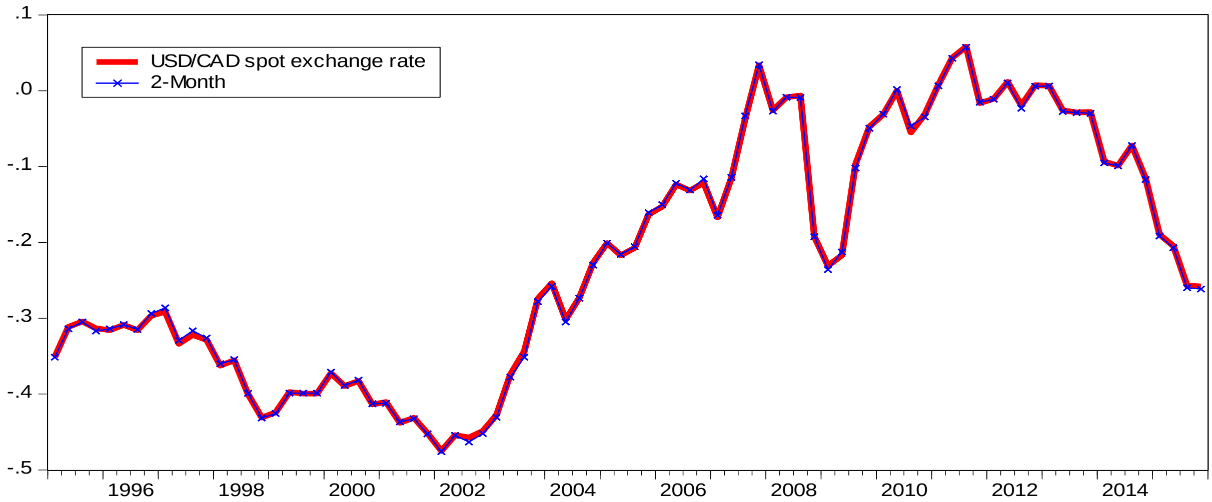

Figure A3.8. The USD/CAD Spot Exchange Rate and Futures Prices 2 Months Prior to Expiry

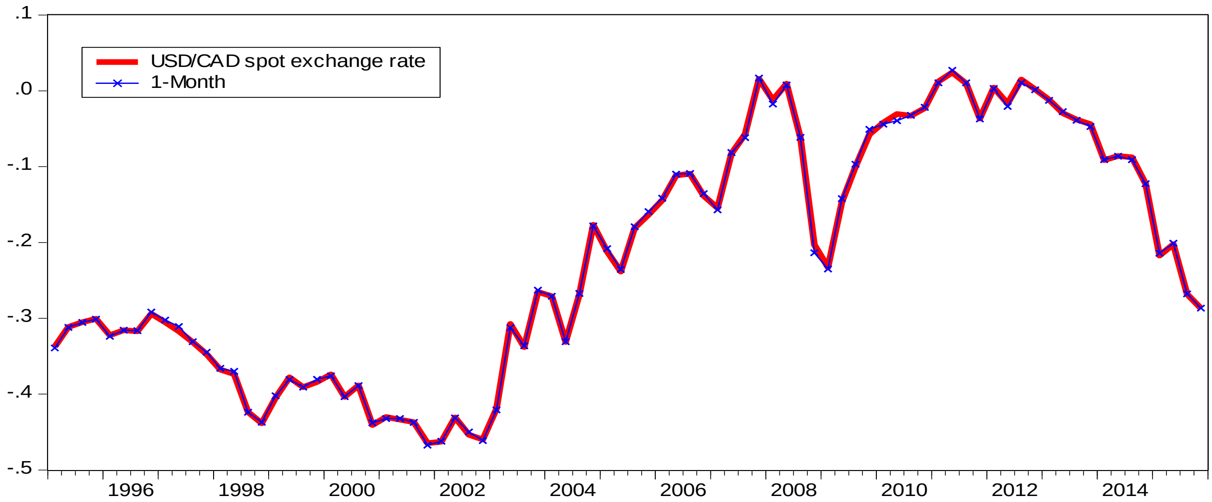

Figure A3.9. The USD/CAD Spot Exchange Rate and Futures Prices 1 Months Prior to Expiry

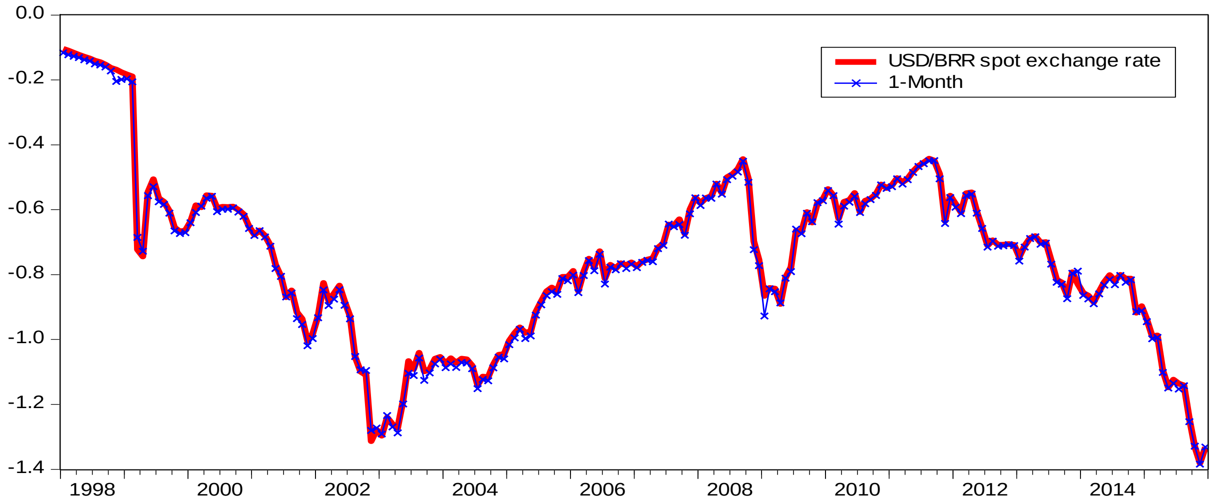

Figure A3.10. The USD/BRL Spot Exchange Rate and Futures Prices 1 Month Prior to Expiry

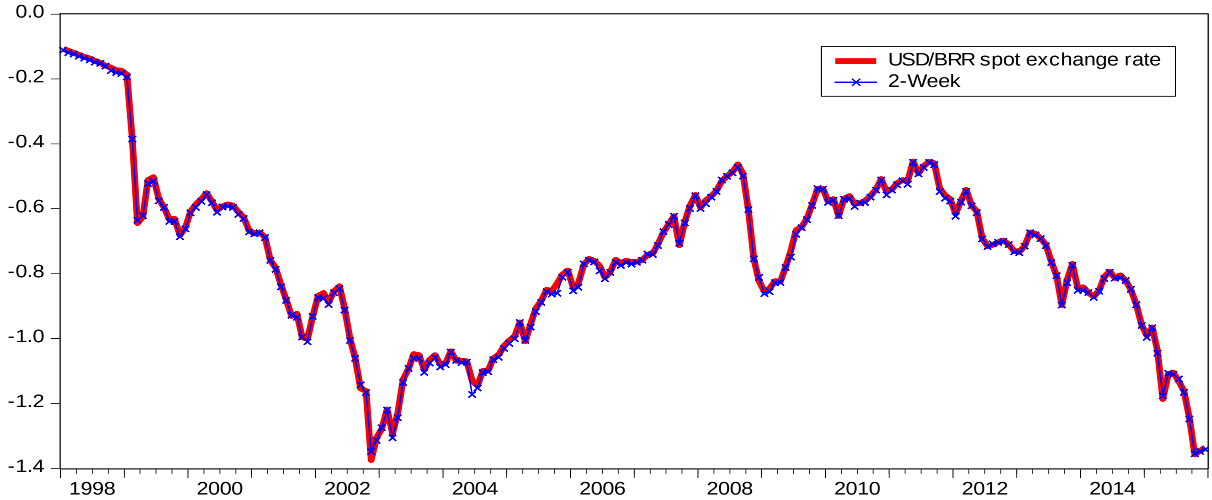

Figure A3.11. The USD/BRL Spot Exchange Rate and Futures Prices 2 Weeks Prior to Expiry

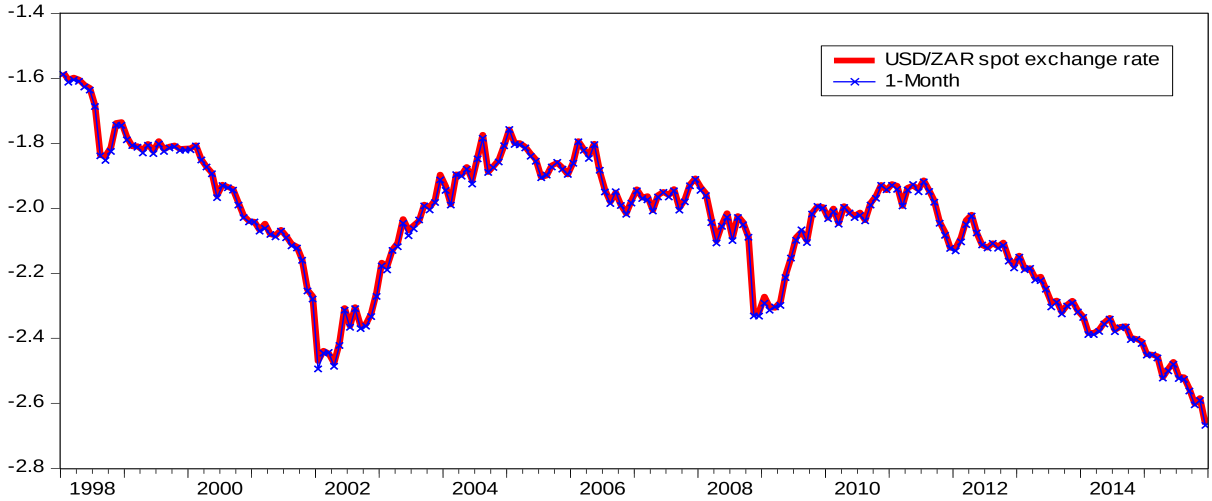

Figure A3.12. The USD/ZAR Spot Exchange Rate and Futures Prices 1 Month Prior to Expiry

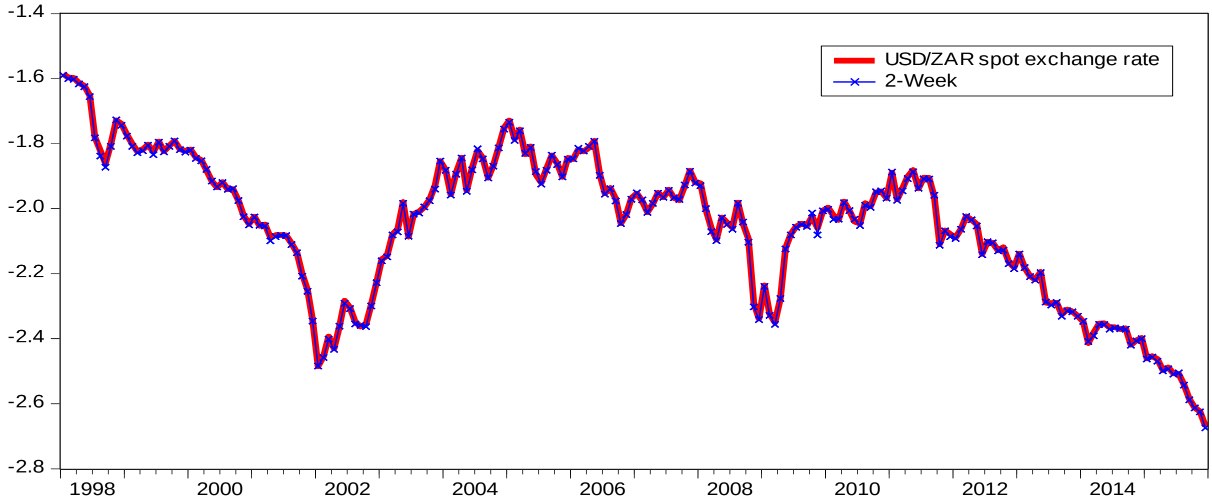

Figure A3.13. The USD/ZAR Spot Exchange Rate and Futures Prices 2 Weeks Prior to Expiry

Appendix A3.4.

Cointegrating Vectors for the USD/GBP futures contracts (Figures A3.14.-A3.16.)

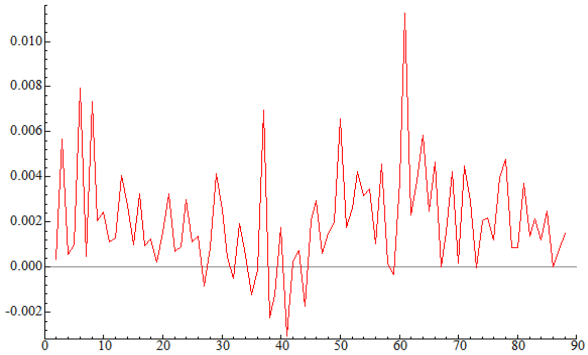

Figure. A3.14. 1-Month Cointegrating Vector

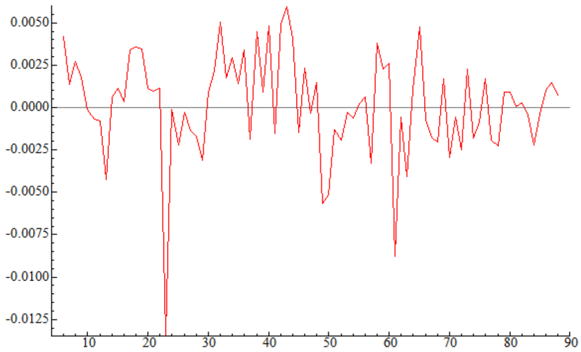

Figure. A3.15. 2-Month Cointegrating Vector

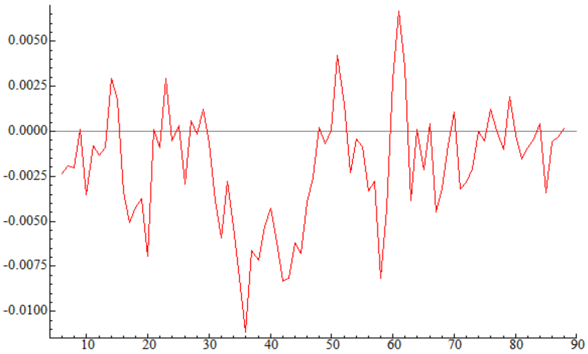

Figure. A3.16. 3-Month Cointegrating Vector

Cointegrating Vectors for the USD/AUD futures contracts (Figures A3.17.-A3.19.)

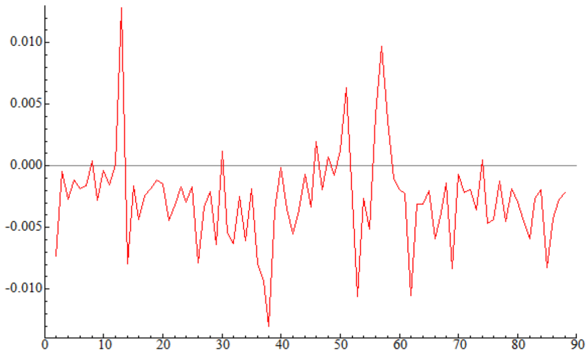

Figure. A3.17. 1-Month Cointegrating Vector

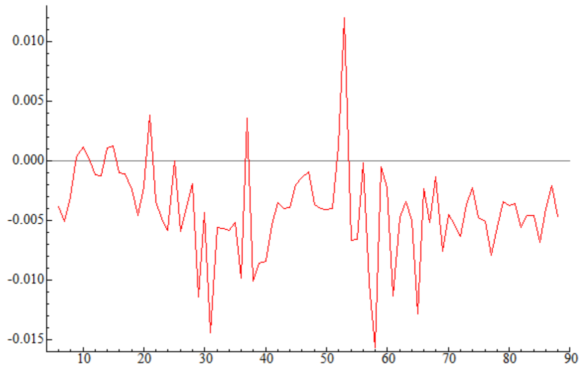

Figure. A3.18. 2-Month Cointegrating Vector

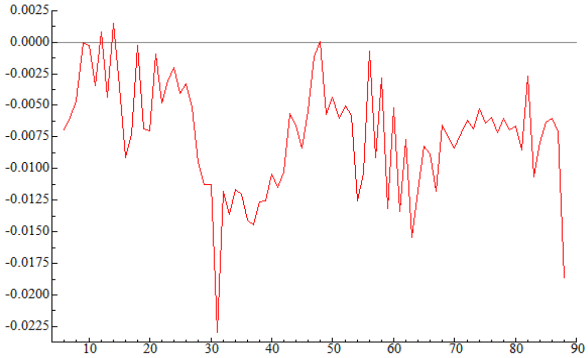

Figure. A3.19. 3-Month Cointegrating Vector

Cointegrating Vectors for the USD/CAD futures contracts (Figures A3.20.-A3.22.)

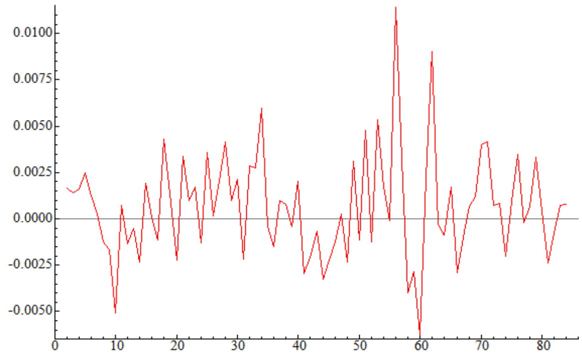

Figure. A3.20. 1-Month Cointegrating Vector

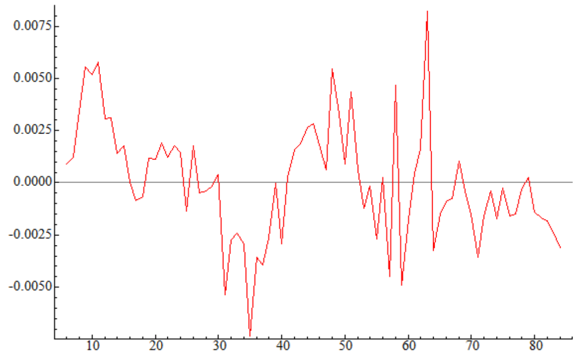

Figure. A3.21. 2-Month Cointegrating Vector

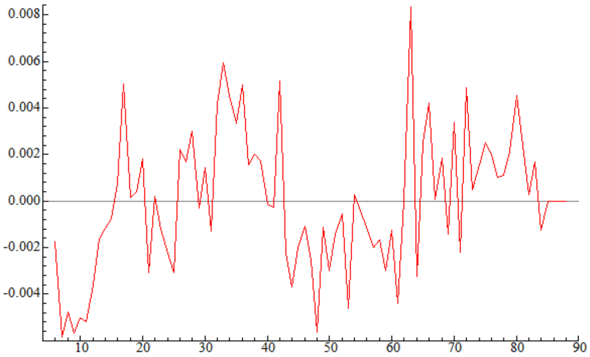

Figure. A3.22. 3-Month Cointegrating Vector

Cointegrating Vectors for the USD/BRL futures contracts (Figures A3.23.-A3.25.)

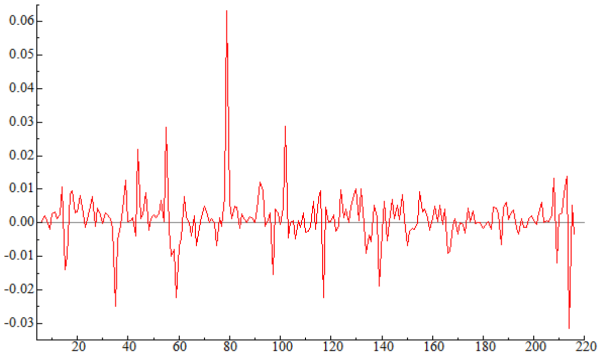

Figure. A3.23. 1-Week Cointegrating Vector

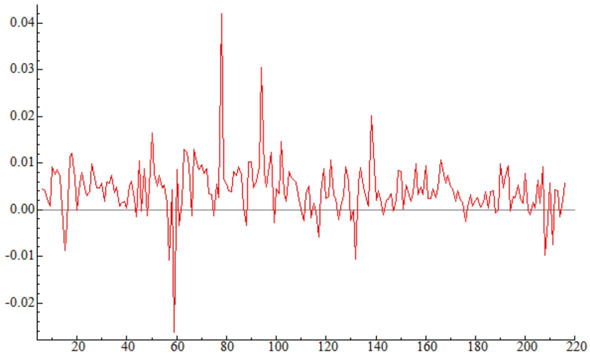

Figure. A3.24. 2-Week Cointegrating Vector

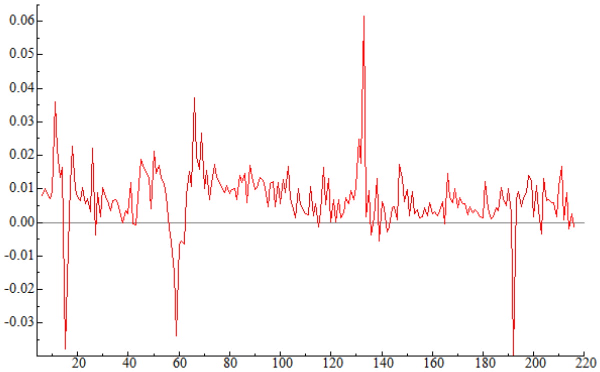

Figure. A3.25. 1-Month Cointegrating Vector

Cointegrating Vectors for the USD/ZAR futures contracts Figures A3.26.-A3.28.)

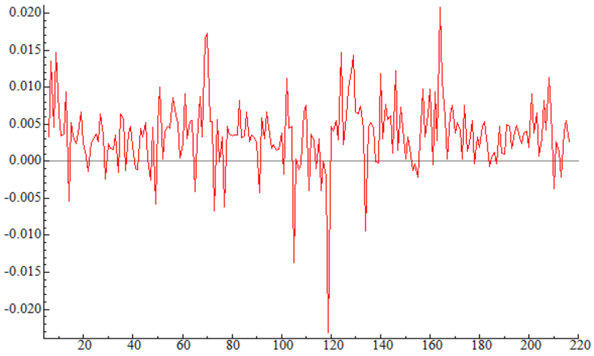

Figure. A3.26. 1-Week Cointegrating Vector

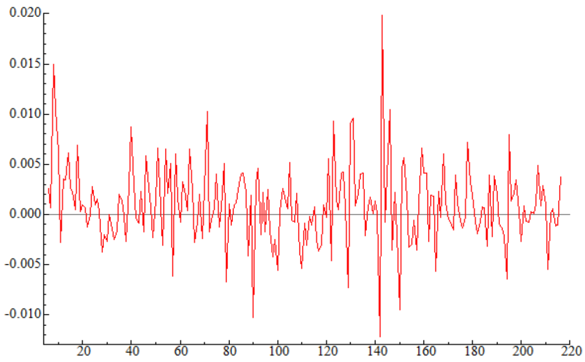

Figure. A3.27. 2-Week Cointegrating Vector

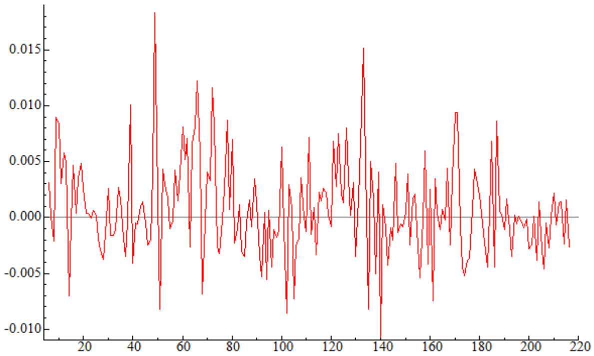

Figure. A3.28. 1-Month Cointegrating Vector

Appendix A3.5.

Figures A3.29. - A3.33. Graphical representation of spot exchange rates

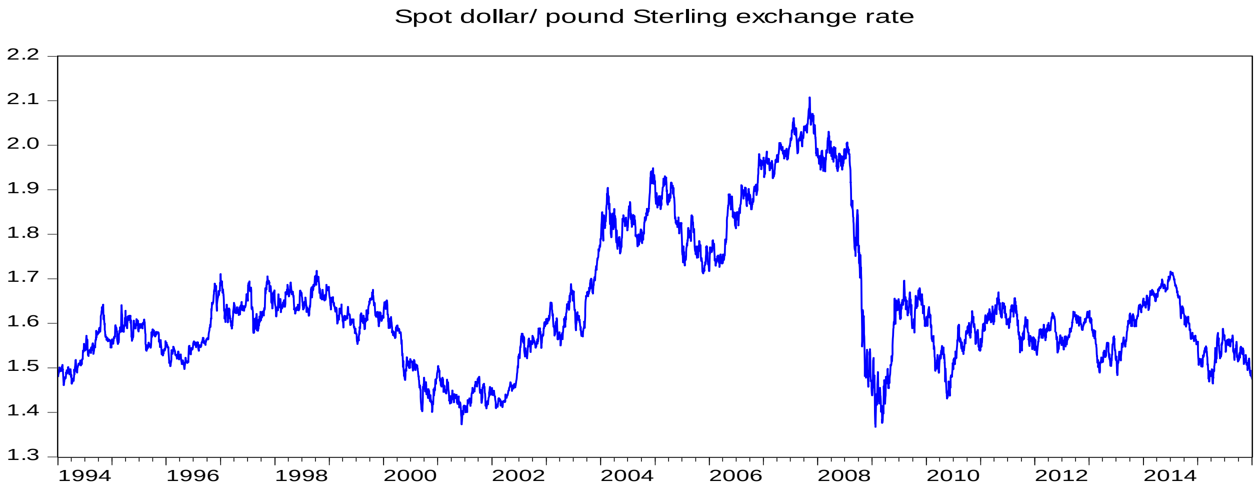

Figure A3.29. Spot USD/GBP Exchange Rate

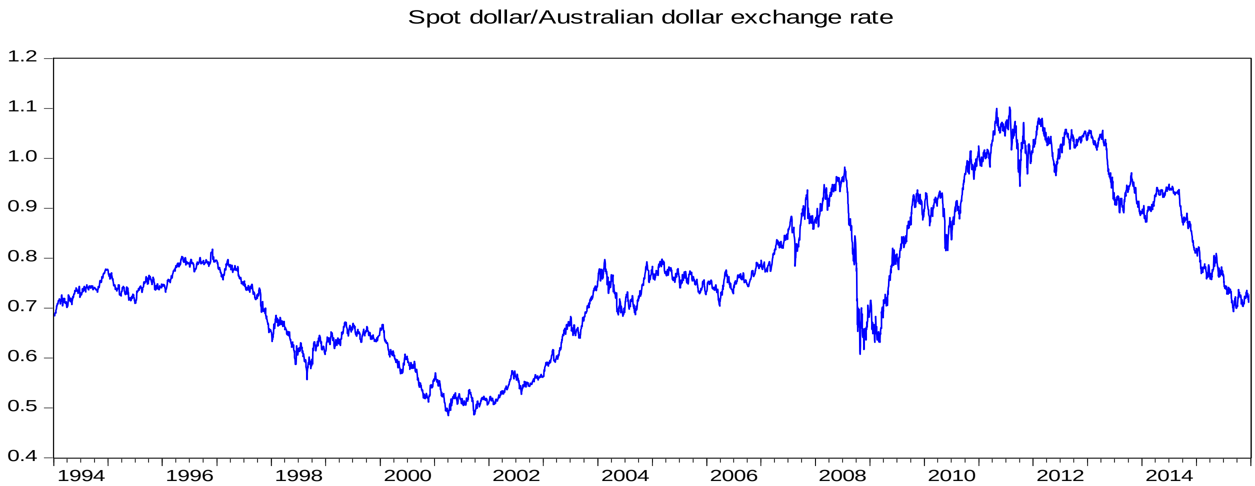

Figure A3.30. Spot USD/Australian Dollar Exchange Rate

Figure A3.31. Spot USD/CAD Exchange Rate

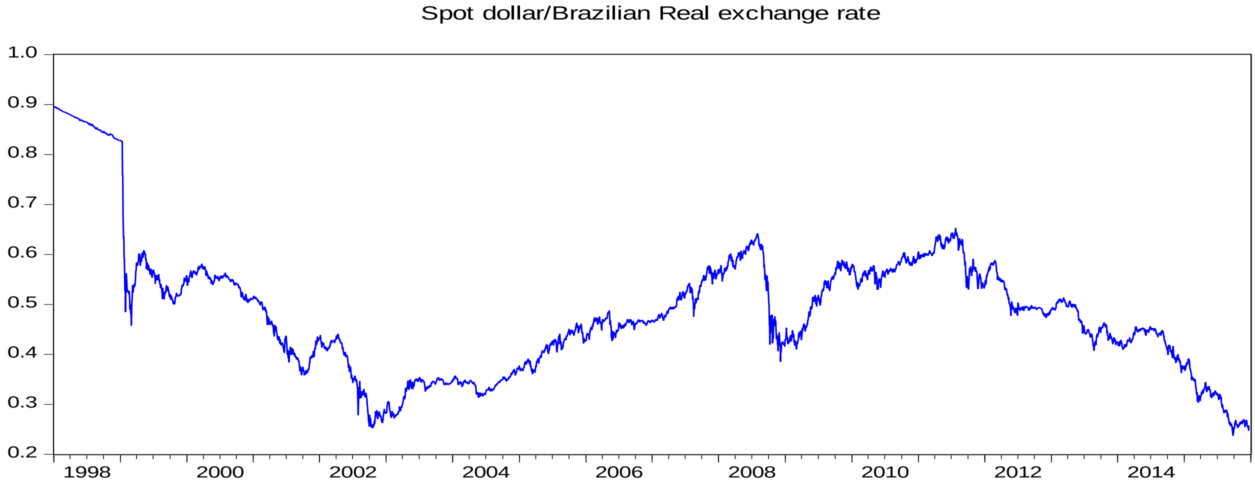

Figure A3.32. Spot USD/Brazilian Real Exchange Rate

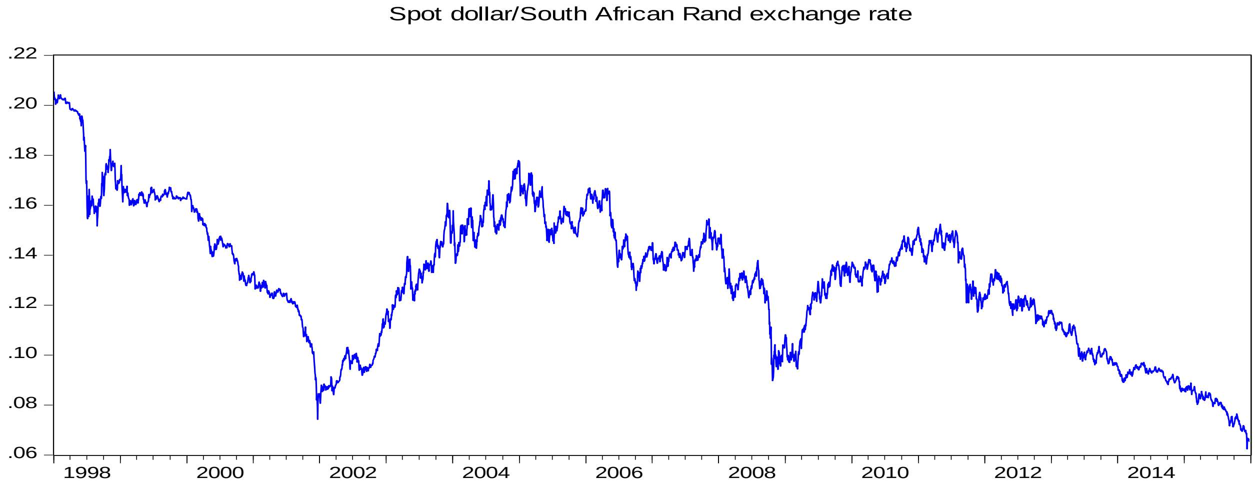

Figure A3.33. Spot USD/South African Rand Exchange Rate

Appendix A3.6.

A correction is applied to each cross-section variance as:

$$\sigma_i^{*2}(w) = \gamma_o + 2\sum_{s=1}^{T-1} k(w)\gamma_s \qquad \text{(A3.1)}$$

Where, $\gamma_0 = \sigma_i^{*2}$, *w=s/l+1* is the bandwidth, *l* is lag truncation in the covariance weighting. The autocovariance can therefore be expressed as:

$$\gamma_s = \frac{1}{T}\sum_{t=s+1}^{T} e_{it}e_{it-s}. \qquad \text{(A3.2.)}$$

The truncated (T) is: $k_T(w) = \begin{Bmatrix} \text{1 for w<1} \\ \text{0 otherwise} \end{Bmatrix}$ (A3.3.)

The Hadri (2000) panel unit root test stipulates a consistent estimate of the residual spectrum at zero frequency. The estimator for the residual spectrum at zero frequency that Hadri (2000) has suggested using, is the Quadratic-Spectral (QS) estimator that is a kernel-based sum-of-covariances estimator is represented as:

$$k_{QS}(w) = \frac{25}{12\pi^2 x^2}\left\{\frac{\sin(6\pi w/5)}{6\pi w/5} \text{-}\cos(6\pi w/5)\right\} \qquad \text{(A3.4.)}$$

Hadri (2000) also presents the Bartlett (BT) and Tukey-Hanning (TH) kernel based estimators for comparison. The LM statistic is not chi-squared, but the following correction follows a standard normal distribution in the limit: $Z_u = \frac{\sqrt{N}(LM_u - \xi_u)}{\zeta_u} \rightarrow N(0,1)$ (A3.5.)

Hadri (2000) shows that ξ_u=1/6 and ζ_u^2 = 1/45. Also for T≥50, the empirical size of the test is approximately 0.0531 and for υ in the range *[100,∞]* the test has maximum power. Expressions for the corrections and estimators, illustrated above, as well as the test statistics presented in Table A3.16., below, are adopted from Beirne et al. (2007).

Table A3.16. Non-parametric correction to the Hadri (2000) test based on alternative kernels	
Kernel estimator	**Test statistic**
QS	0.925401
Truncated	0.935337
TH	1.297038
Bartlett	1.662121

We note that the test is for a one-sided inference, which for a test at the 5% level implies a critical value of 1.645. When the tests are ordered by speed of decay, the test statistics based on TH, QS and T kernels all accept the null of stationarity, while the test using the Bartlett kernel marginally fails at the 5% level. From Table A3.16. we observe that the BT kernel rejects the null (t-critical > 1.66) as it may not capture the autocorrelation in the data when it is highly autoregressive. More recent research might suggest using a lag truncation of T/3=22 (Beirne et al. 2007).

CHAPTER FOUR
TRADED GOODS CONSUMPTION SMOOTHING AND THE US DOLLAR VIS-À-VIS THE JAPANESE YEN

4.1. Introduction

The near random walk behaviour exhibited by real exchange rates is generally substantiated by the random walk behaviour of underlying elements. In this chapter a monetary model of the USD/JPY exchange rate is proposed whereby variables employed in the fixed-factor neoclassical model (Obstfeld, 1982; Frenkel and Razin, 1986; Frankel 1987b) are merged together with those employed in the standard RID monetary model formulated by Frankel (1979a) and share prices (Friedman, 1988) to create a hybrid monetary model akin to that of Hunter and Ali (2014). Monetary models are based on the link between exchange rates and prices that change together domestically and abroad. These models have historically failed to identify a relationship between exchange rates and prices as a result of unstable money demand functions and the breakdown of PPP.

This study pertains to traditional macroeconomics and we draw upon the standard monetary model, finance, pricing, asset pricing theory as well as dynamic micro-foundation-based models of the real exchange rate (Obstfeld, 1982). The conventional economics system is developed to better assess endogeneity and examine cointegration as a long-run explanation of the exchange rate as part of a system, distant from a random walk as it is non-stationary. A model of the real USD/JPY exchange rate using the instrumental variables considered by Rogoff (1992) is first estimated and a test of cointegration is carried out by employing the Johansen (1995) methodology as compared to the Engle-Granger (1987) two-step error correction model. Cointegration is subsequently examined amongst variables related to the monetary base and the USD/JPY exchange rate in the standard RID model developed by Frankel (1979a). To complete the analysis the standard USD/JPY RID model

(Frankel, 1979a) is merged with the instrumental variables employed by Rogoff (1992) and real share prices (Friedman, 1988). The behaviour underlying the USD/JPY exchange rate, based on stable money demand and PPP, is expected to be driven by non-stationary prices.

We contribute to the extant literature by showing that the conventional RID monetary model of the USD/JPY exchange rate for an extended data set in a period characterized by extraordinary macro policy initiatives can be rejected. We also show that the USD/JPY exchange rate is not readily distinguished from a random walk in the context of a monetary model that considers traded and non-traded goods productivity differentials.

4.1.2. The Exchange Rate and Economic Fundamentals of the United States and Japan

The Balassa-Samuelson effect (Balassa, 1964 and Samuelson, 1964) describes the relationship between exchange rates and cross-country differences in the price of non-traded goods whereby an overvalued domestic currency hinders domestic employment and salary growth so restricts the current account and consequently private investment as well as government spending abroad. The neoclassical model formulated by Rogoff (1992) considers the Balassa-Samuelson effect (Balassa, 1964 and Samuelson, 1964) and assumes that trivial shocks to non-traded goods productivity affect consumption smoothing in the traded as well as the non-traded goods sectors. The model incorporates variables related to productivity in manufacturing, government consumption expenditure (Froot and Rogoff, 1991) and the real oil price.

Rogoff’s (1992) study of consumption smoothing and the USD/JPY real exchange rate is based on data from 1975:Q1-1990:Q3, a period characterized by high inflation. The study assumes rational expectations and employs instrumental variables in a conventional systems equation (Engle and Granger, 1987), as compared to a VAR methodology, that does not consider endogeneity since information is assumed to be predetermined. Meese and

Rogoff (1983a) examine the performance of a similar systems equation employing variables related to the monetary base and find that such a model lacks the empirical ability to beat the out-of-sample naïve random walk. In this analysis the structural consumption smoothing model proposed by Rogoff (1992) is revised to account for the current period of zero lower-bound interest rates and subsequent economic developments by employing a dynamic structural equation and testing for cointegration.

Hunter and Ali (2014) reject the proposition that the exchange rate follows a random walk in a hybrid monetary model of the JPY/USA nominal exchange rate. The exchange rate in itself follows a stochastic trend but in the context of a system it is cointegrated and the cointegrating relationships can be defined as moving averages in addition to the random walk behaviour of the exchange rate. We are therefore able to state that the error correction type term in the structural model captures features that can be described as moving averages. It would therefore not be inconsistent and certainly possible that a model exhibiting purely cointegration and no further dynamics of the exchange rate would resonate with Rogoff's (1992) model as further correlation is captured by the error correction term. A single variable is therefore able to retain the behaviour of all the other level relationships within the theoretically related variables. Therefore, if random walk behaviour is detected in the context of a single variable one would employ instrumental variables and consider decomposition of the error term in a structural framework in an alternative way. By reference to arbitrage pricing theory the exchange rate is perceived to be an asset price that follows a random walk and this framework encompasses factors that are driven by new information emanating from the macro-economy, amongst other things. We perceive from Figure 4.1. that the real USD/JPY exchange rates, in levels, deflated by CPI follows a random walk.

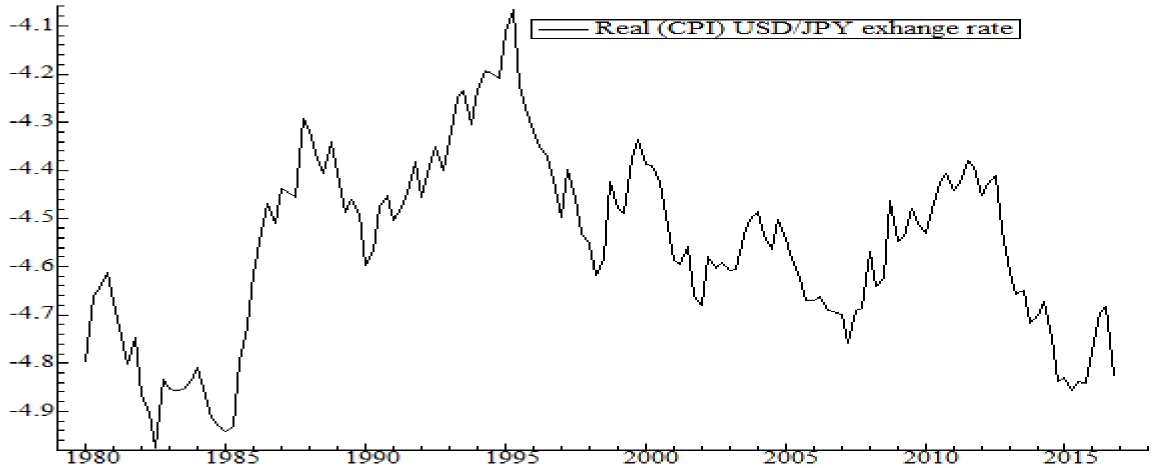

Figure 4.1. The Real USD/JPY Exchange Rate, in Levels, Deflated by CPI

Hunter and Ali (2014) in their hybrid model of the USD/JPY exchange rate find a single cointegrating relation by following the traded goods consumption model (Rogoff, 1992) as filtered through Pilbeam (2013) and incorporate a variable related to government consumption expenditure because of inefficiencies related to the non-traded sector linked to government where government activity is non-traded. The relative price of traded goods will however satisfy the PPP condition. We can explain consumption as a random walk (Hall, 1978) however it has further components that drive the random walk. Figure 4.2. illustrates the relationship between relative government spending and the real USD/JPY exchange rate. We perceive that when government spending in the USA is greater than in Japan the JPY depreciates against the USD (see Appendix A4.1. Figure A4.1. for the USD/JPY against nominal Japanese government spending (GS), Figure A4.2. for the USD/JPY against real Japanese GS and Figure A4.3. for the USD/JPY against the nominal GS differential).

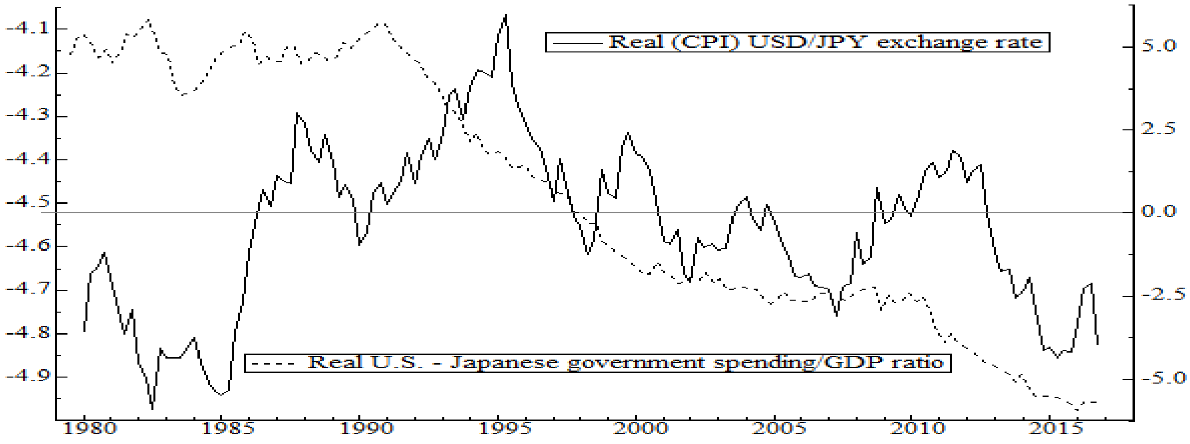

Figure 4.2. The Real (CPI) USD/JPY Exchange Rate (Left-Hand Scale) Versus the Real US - Japanese Government Spending/GDP Ratio (Right-Hand Scale)

The JPY is historically overvalued (Obstfeld, 2009) and persists to be overvalued so PPP does not hold otherwise there would be no overvaluation. Summers and Heston (1991) reconstructed price series to pay attention to overvaluation of the JPY so that the relative price of traded goods will satisfy PPP. Obstfeld (1982) develops a stochastic model of the real exchange rate and tackles issues such as traded and non-traded sectors productivity that are trend stationary. However, he argues that some series are most likely predicted by a trend so the data is not captured in a particularly effective way.

Monetary models of the USD/JPY exchange rate have especially failed to produce a meaningful long-run explanation based on unstable money demand functions (Hendry and Ericsson, 1991; Caporale and Pittis, 2001; Chinn and Moore, 2011) and the collapse of PPP (Sarno and Taylor, 2002). Collapse of PPP has been ascribed to productivity differentials. Wang and Dunne (2003) reproduce a monetary model of the USD/JPY exchange rate for the period 1973:Q1- 1996:Q4 and report that differentials related to productivity and government spending as well as the real oil price are responsible for exchange rate volatility. The invalidity of PPP is evident in Figure 4.3. where relative national prices are incapable of predicting the spot USD/JPY exchange rate.

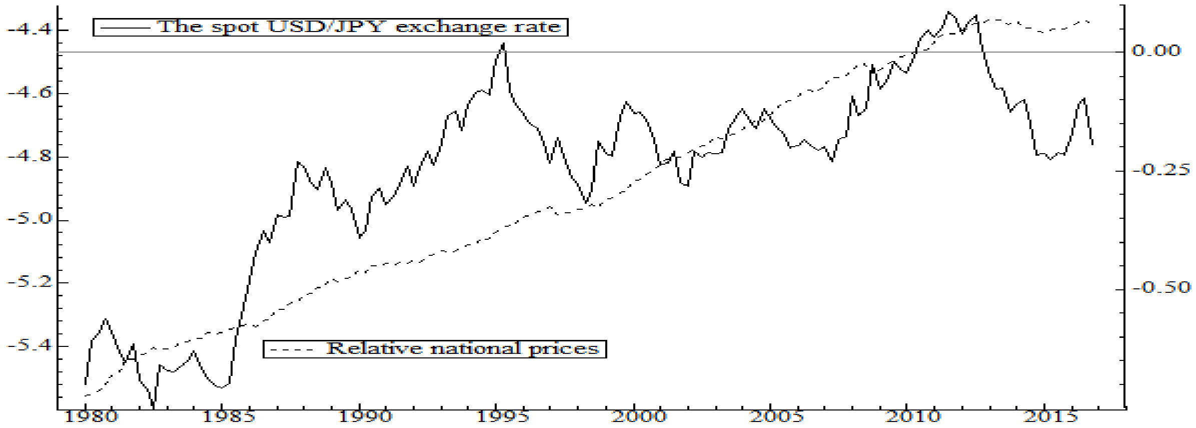

Figure 4.3. The USD/JPY Nominal Exchange Rate (Left-Hand Scale) Versus Relative Consumer Prices (Right-Hand Scale)

Friedman (1988) and many others (e.g. McCornac, 1991 and Caruso, 2001) enhance stability of money demand function in the conventional monetary model by including real share prices that consider real economic activity and long-run expectations. Other incentive for including real share prices in the monetary model is that financial market experts and the media perceive a link between exchange rates and share prices (Phylaktis and Ravazzola, 2005; Caporale et al., 2014). Figure 4.4. illustrates a positive correlation between the USD/JPY exchange rate and relative real share prices.

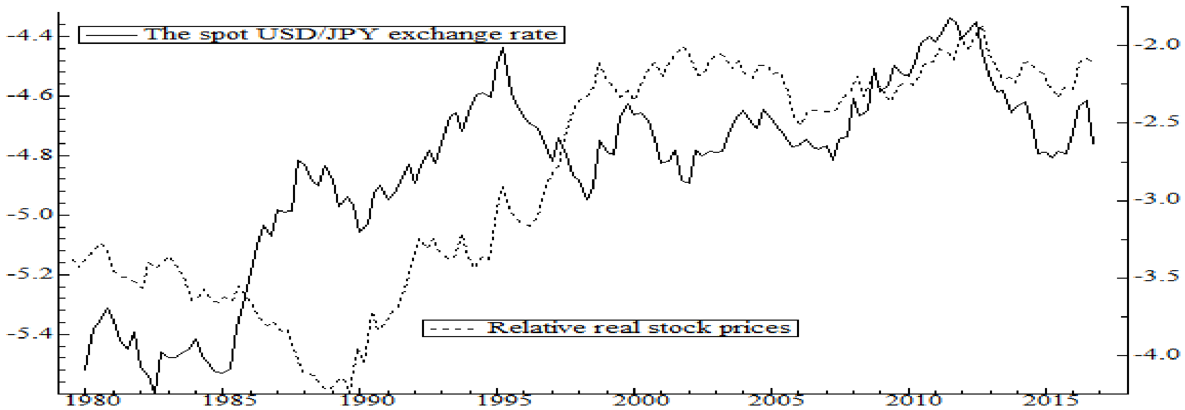

Figure 4.4. The USD/JPY Nominal Exchange Rate (Left-Hand Scale) Versus Relative Real Stock Prices (Right-Hand Scale)

Rogoff (1992) constructs an exchange rate model that incorporates underlying shocks to productivity and the Euler condition, akin to the capital asset pricing model but referring to consumption and not assets, and explains consumption as a random walk (Hall, 1978). However, he indicates that further components exist, and the random walk is driven by these extra factors. Hsieh (1982) examines departures of the JPY/USD exchange rate from PPP by considering labour productivity differentials between the traded and non-traded goods sectors and finds that time series analysis produces a more constructive validation of the relative differential model than cross-section regressions.

Exchange rates have in addition shown to be responsive to supply shocks linked to the real price of oil (Johansen and Juselius, 1992). Amano and van Norden (1998) illustrate that higher oil prices lead to a re-valuation of the USD since oil prices are quoted in terms of the USD. Figure 4.5. illustrates a positive correlation between the real oil price and real USD/JPY exchange rate deflated by CPI.

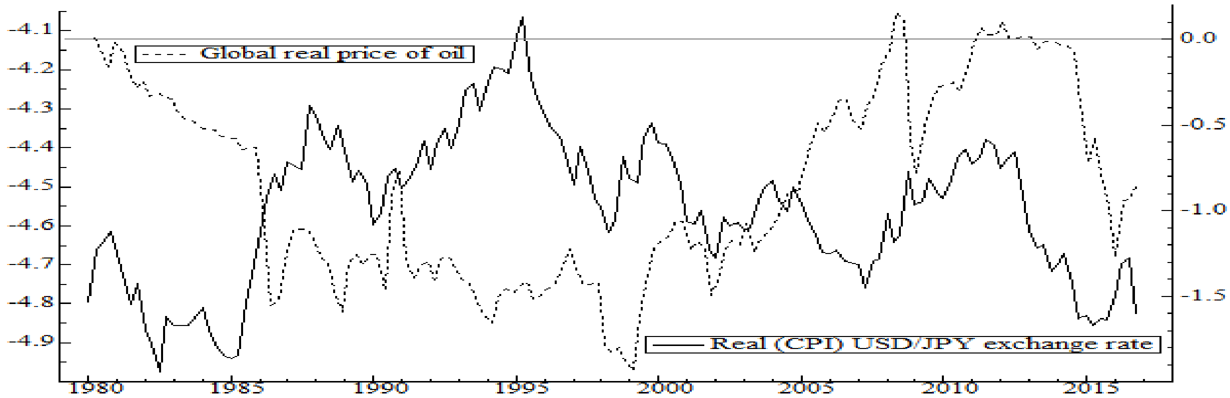

Figure 4.5. The Real (CPI) USD/JPY Exchange Rate (Left-Hand Scale) Versus the Global Real Price of Oil (Right-Hand Scale)

Chinn (1997a) investigates both long and short-run determinants of the real exchange rate using data from 1970 to 1992 and argues that average labour productivity is reduced by the impact of demand e.g. amassing labour so an advantage of reproducing the analysis by means of cointegration avoids such problems so long as the cyclical element of labour productivity is stationary. In another study Chinn (1997b) employs quarterly data from 1974 to 1993 and shows that the traded goods productivity differential is non-stationary while the differential related to non-traded goods productivity is stationary. So, he employs a test of cointegration that allows for a trend in the data generating procedure.

Mckinnon and Ohno (1997) discuss Japan’s sluggish economy and provide an explanation as to why aggregate private demand has not recouped following collapse of Japan’s stock and real estate markets in 1991 and why deflationary economic trends persist. Mckinnon and Ohno (1997) suggest that trading conflict with the USA and the discrepancy between Japan’s current account surpluses and the United States’ current account deficit is responsible for the Japanese Yens’ overvaluation and low Japanese wholesale prices as well as interest rates that are considerably lower on Japanese assets than on American assets.

Hamada and Okada (2009) demonstrate that fiscal and global factors as well as non-fiscal and domestic elements had an impact on Japans’ stagnation. The Plaza Agreement of 1985 as well as monetary exchange rate rule served to strengthen the JPY that only debilitated Japanese industry and caused a decay in productivity and development. Figure 4.6. illustrates that manufacturing productivity in Japan has not risen to pre-2007-2009 financial crisis levels and Figure 4.7. illustrates that GDP in Japan has experienced a severe decline since the 2008 financial crisis and as compared to the USA is experiencing slow growth as pre-crisis levels were only just reached in 2013 (see Appendix A4.1. Figure A4.4. for the real USD/JPY exchange rate against Japanese labour productivity in manufacturing).

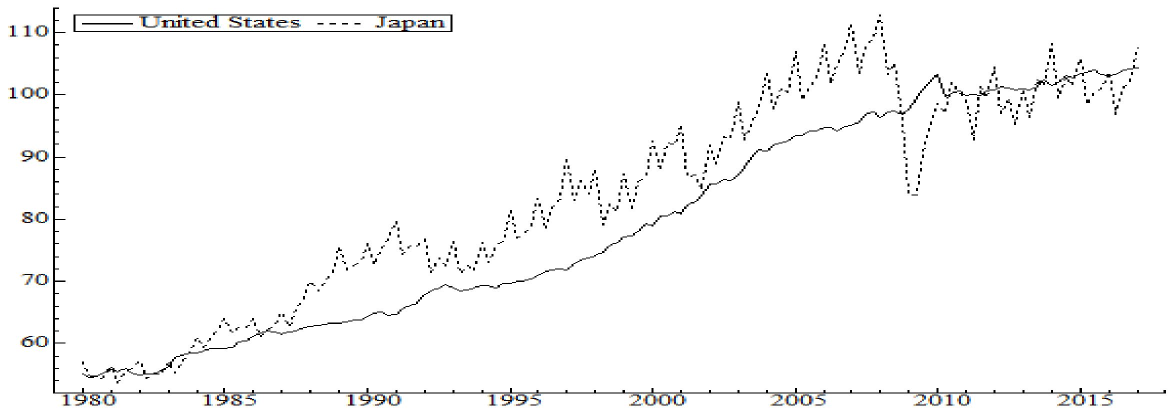

Figure 4.6. Productivity in Manufacturing - the USA Versus Japan

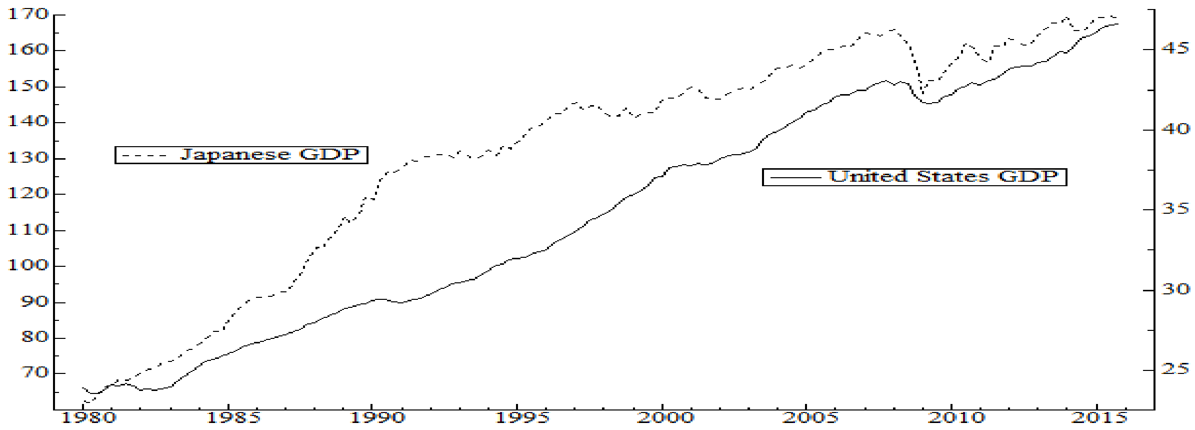

Figure 4.7. Gross Domestic Product – the USA Versus Japan

An and Kim (2010) employ monthly data from 1970 to 2005 in a structural VAR and examine implications of nominal and real shocks to the USD/JPY exchange rate by imposing restrictions in both the short and long-run. An and Kim (2010) show that exchange rate volatility stems from shocks related to the exchange rate, the real oil price, the productivity differential, relative demand and relative monetary policy and that the exchange rate appreciates in response to positive asymmetric supply (demand) shocks that are shown to be responsible for a major part of exchange rate volatility.

The rest of this chapter is ordered as follows. Section 4.2 briefly reviews monetary models of the exchange rate including the intertemporal model (Rogoff, 1992), the standard RID model and the adjusted hybrid monetary model. Section 4.3 describes the econometric approach the data employed in this analysis. Empirical results are reported in Section 4.4 and concluding comments are presented in Section 4.5.

4.2. The Monetary Models of Exchange Rate Determination

4.2.1. The Intertemporal Model (Rogoff, 1992)

This model distinguishes between traded and nontraded goods so is applied to further understand the impact of productivity and government consumption expenditure shocks. This model resembles that run by Dornbusch (1983) and Froot and Rogoff (1991) amid others. The model assumes a small open economy that has little influence over the real global interest rate.

Production relates to traded (T) and nontraded (N) goods and functions representative of each sector are as follows:

$$Y_{Tt} = A_{Tt} L_{Tt}^{\theta T} K_{Tt}^{1-\theta T} \tag{4.1.}$$

$$Y_{Nt} = A_{Nt} L_{Nt}^{\theta N} K_{Nt}^{1-\theta N} \tag{4.2.}$$

Where Y_T represents output in the traded sector, Y_N represents output in the non-traded sector, A_I, K_I, and L_I represent stochastic productivity, capital and labour shocks in the sectors for traded and non-traded goods I *where* I=(T,N). It is assumed that capital and labour are not interchangeable between sectors within the economy.

A time-divisible utility function that describes an economy populated by a representative agent is illustrated as follows:

$$V_t = E_t \sum_{s=0}^{\infty} \beta^{s-t} \left[\frac{(C_{Ns}^{\alpha} C_{Ts}^{1-\alpha})^{1-\gamma}}{1-\gamma} \right] \tag{4.3.}$$

Where *E* represents the expectations operator, β represents the relative discount rate, and C_{It} is the consumption of good *I* during period t; γ is the reverse elasticity of the intertemporal swap.

The government and private sectors have unrestricted access to world financial markets where noncontingent bonds are dealt at a gross interest rate *R*, calculated in term of traded goods. Trading is conducted under the assumption that shocks to traded goods productivity are not smoothed by diversification. Ricardian equivalence holds so that government consumption expenditure remains unchanged as the public sets aside money surplus funds to pay future taxes. Subsequently the government’s fiscal plan and that of the general public can be consolidated so that the individual's intertemporal fiscal limit can be presented as:

$$F_{t+1} = R(F_t + Z_{Tt} - C_{Tt} - P_t C_{Nt} - G_{Tt} - P_t G_{Nt}) \tag{4.4.}$$

where F_t is the general individual's holdings of equity abroad beginning period *t*, *R* the gross global real interest rate (computed in tradeables), *P* is the relative price of non-traded goods in terms of traded goods, and Z_t represents overall earnings from productivity at home (computed in terms tradeables). G_t represents government consumption of tradeables, and G_{Nt} is government consumption of nontraded goods. In this instance it is assumed that government consumption has no influence on the utility of private consumption (Obstfeld, 1982).

Non-traded goods cannot be swapped intertemporally, so consumption of non-traded goods should correspond to non-traded goods productivity domestically. Since by definition, there is no way to swap nontraded goods intertemporally, domestic consumption of nontraded goods must equal domestic output of nontraded goods every time so that:

$$Y_{Nt} = C_{Nt} + G_{Nt} \tag{4.5}$$

The consumption of traded goods can under the assumption of free capital markets be smoothed over time. The intertemporal fiscal constraint for the entire economy can - under the no Ponzi scheme condition and by employing the intertemporal fiscal limit in (4.4) and the non-traded goods equilibrium condition in equation (4.5.) - be presented as:

$$\sum_{s=0}^{\infty} \frac{C_{Ts}}{R^{s-t}} = F_t + \sum_{s=0}^{\infty} \frac{Y_{Ts} - G_{Ts}}{R^{s-t}} \tag{4.6}$$

The maximization condition at time t suggests that the price of nontraded goods (P_t) is contingent upon the differential between consumption of both traded and non-traded goods and is as follows:

$$P_t = \frac{\alpha C}{(1-\alpha)C_{Nt}} \tag{4.7}$$

The relative price of non-traded goods and the real exchange rate henceforth will be used reciprocally since the CPI deflator is founded on consumption for the input versus the output function over time as represented in equation (4.3). Variation in the relative value of non-traded goods produces the sole cause of real exchange rate changes under the assumption that the relative value of imports and exports or terms of trade remain unchanged.

Since consumption of non-traded goods at any time hinges on the supply at present equations (4.5.) and (4.7.) can be amalgamated to produce:

$$P_t = \frac{\alpha C_{Tt}}{(1-\alpha)(Y_{Nt} - G_{Nt})} \tag{4.8.}$$

The price of non-traded goods increases with heightened government consumption expenditure as the net supply accessible to the private sector is reduced. Consequently, agents smooth traded goods consumption over time so that the resulting marginal utility function is as follows:

$$\left(C_{Nt} / C_{Tt}\right)^{\alpha} \cdot (C_{Nt}^{\alpha} C_{Tt}^{1-\alpha})^{-\gamma} = \beta R E_t (C_{Nt+1} / C_{Tt+1})^{\alpha} \cdot (C_{Nt+1}^{\alpha} C_{Tt+1}^{1-\alpha})^{-\gamma} \tag{4.9.}$$

Before examining the behaviour of real exchange rates implied by the “fixed factor/open capital markets model,” it is helpful to review the classical model of real exchange rates developed by Balassa (1964) and Samuelson (1964). The HBS model, discussed in Chapter 2 (see section 2.2.2), assumes that capital and labour in both the traded and non-traded goods sectors are fixed at home and that streams of both inputs are promptly balanced. Under the assumptions of unrestricted global capital mobility and interchangeable labour across sectors the relative price of non-traded goods in equation

(4.8.) is directed by labour while government consumption expenditure and the individual's fiscal limit exhibit little influence.

Under the assumption of absolute interchangeability between sectors the yield maximization can be illustrated as follows:

$$R = (1-\theta_T)A_T(K_T / L_T)^{-\theta T} = P(1-\theta_N)A_N(K_N / L_N)^{-\theta N} \tag{4.10.}$$

$$W = \theta_T A_T(K_T / L_T)^{1-\theta T} = P\theta_N A_N(K_N / L_N)^{1-\theta N} \tag{4.11.}$$

where W is the price of labour or wage rate. K_T/L_N, is the ratio of capital to labour related to the first equation in (4.10), under the condition of free worldwide capital mobility. While the price of labour is directed by the first equation in (4.11). Equations (4.10) and (4.11) exhibit an effect on K_N/L_N and P.

The traditional Balassa-Samuelson model is derived by logarithmically differentiating equation (4.10) and (4.11) and is illustrated as:

$$dp = (\theta_N / \theta_T)d\alpha_T - d\alpha_N \tag{4.12.}$$

where lowercase letters represent logarithms and *d* represents differentials. The term θ_N/θ_T in equation (4.12.) is represented as such since non-traded goods are comparatively more work concentrated so that an increase in α_t prompts an increase in wages and as a consequence the relative price of non-traded goods is greater than that of traded goods. Equation (4.12.) is sustained even if factors of production are not immediately interchangeable between sectors and if the economy is closed to cross-border borrowing and lending. Since traded goods cannot be traded over time, the marginal utility function for consumption smoothing overtime is substituted for the market clearing function as follows:

$$Y_{Tt} = C_{Tt} + G_{Tt} \tag{4.13}$$

Merger of equations (4.8.) and (4.13) yields:

$$dp \cong \zeta_T da_T - \zeta_N da_N - [(\zeta_{T-1})dg_T - (\zeta_{N-1})dg_N] \tag{4.14}$$

where ζ_I is the output to consumption ratio in sector *I*. Shocks to productivity have an inverse impact to that of the Balassa-Samuelson model. Cumulative shocks to government spending are however of consequence in the fixed factor neoclassical model. In equation (4.14) 'p' has a tendency to increase with an increase in the consumption of non-traded goods when consumption in the government sector is greater than that in the private sector. The open capital markets model is considered next to correct for discrepancy between expected and unexpected variations.

The assumptions intrinsic to equation (4.14.) are a more accurate estimate of the real world than the assumptions behind equation (4.12.). Testimonial to the Balassa (1964) and Samuelson (1964) effect is evident in economies where capital markets are restricted, and labour is inflexible as was the instance for Japan during the 1950's and 1960's.

The real exchange rate, in the event that shocks to productivity are homoscedastic and factors are distinct to each sector, exhibits dissimilar behaviour to the Balassa-Samuelson model or the fixed capital markets model. Under the assumption that the latent shocks to productivity (the A_j's) are homoscedastic the Euler condition intrinsic to equation (4.9.) can be estimated on the basis that Rβ=1 as follows:

$$E_t(c_{Tt+1} - c_{Tt}) \cong \frac{\alpha(1-\gamma)}{\gamma + \alpha(1-\gamma)} E_t(C_{Nt+1} - C_{Nt}) \tag{4.15}$$

In the event that the consumption of traded goods is rigid then (4.15) reduces to a logarithmic estimation to the random walk traded goods consumption model developed by Hall (1978). The real exchange rate is therefore affected and by taking logarithms on either side of equation (4.7.) one derives $p_t = c_{Tt} - c_{Nt} + \log(\alpha / (1-\alpha))$ and consequently:

$$P_{t+1} - P_t = (c_{Tt+1} - c_{Nt+1}) - (c_{Tt} - c_{Nt}) \tag{4.16.}$$

By joint consideration of equations (4.15.) and (4.16.) and ruling out shocks to private consumption of non-traded goods, the log real exchange rate is anticipated to follow a random walk notwithstanding autocorrelation in traded goods productivity shocks. Similarly, a trend in the real exchange rate may not be evident if a trend is present in productivity growth.

The real exchange rate is expected to exhibit greater unpredictability in the Balassa-Samuelson model than in the fixed-factor neoclassical model in addition expected and unexpected shocks to traded goods productivity exhibit little influence in the fixed factor model but in the Balassa-Samuelson model are of equal significance and are expected to exert a similar impact on the real exchange rate.

Factors are conditional on the cost of adjustment and are assumed to be distinct to each sector in the short to medium term with numerous factors expected to be flexible in the long-term. Real exchange rate variation is expected to be persistent despite the impermanence of shocks to traded goods productivity. The model is therefore more suitable for analysis at short intervals as compared to extended intervals over a few years.

The log-linear model is derived according to the Euler condition in equation (4.9.) and under the assumption that shocks to productivity in the traded and non-traded goods sectors are lognormally distributed with homoscedastic error terms: [After normalising for expediency Y=A in both sectors.]

$$\alpha_{Nt+1} = \varphi \alpha_{Nt} + \varepsilon_{Nt} \tag{4.17}$$

$$\alpha_{Tt+1} = \rho \alpha_{Tt} + \varepsilon_{Tt} \tag{4.18}$$

where 0≤φ, ρ≥1, and the ε's are chosen from error terms that are independent over time and sectors. In the instance where φ=1 the real interest rate remains steady if the consumption of non-traded goods behaves like a random walk and according to equation (4.15) consumption of traded goods is not inclined to shift.

The fiscal constraint, in logarithms to a first order estimation, in equation (4.6) is therefore upheld when a trend in the consumption of traded goods is absent. Under the assumption that φ=1 and by employing equation (4.15.) and (4.18.) one derives:

$$c_{Tt+1} - c_{Tt} = \frac{R-1}{R-\rho}(\alpha_{Tt+1} - \rho\alpha_{Tt}) \tag{4.19.}$$

where for expediency G_T has been normalized to zero. Unexpected fluctuation in traded goods proceeds over the span of a lifetime affect variation to consumption provided that nontraded goods consumption behaves like a random walk. When φ=1, equations (4.16).,

(4.17.) and (4.19) suggest:

$$p_{t+1} - p_t = \frac{R-1}{R-\rho}(\alpha_{Tt+1} - \rho\alpha_{Tt}) - (\alpha_{Nt+1} - \alpha_{Nt}) \tag{4.20.}$$

Since government expenditure behaves like a random walk and is predominantly a non-traded goods equation (4.19.) turns into (4.21.), an essential function of the model:

$$P_{t+1} - P_t = \frac{R-1}{R-\rho}(\alpha_{Tt+1} - \rho\alpha_{Tt}) - \zeta_N(\alpha_{Nt+1} - \alpha_{Nt}) + (\zeta_N - 1)(g_{Nt+1} - g_{Nt}) \tag{4.21.}$$

where ζ_N is the ratio of nontraded goods output to nontraded goods consumption.

In the event that shocks to non-traded goods are transient one may let φ<1. In this instance, prices are instantly impacted via equation (4.7) while shocks to non-traded goods productivity also exhibit an impact on the real consumption-based interest rate and so expectedly and unexpectedly shift the direction of traded goods consumption which influences real exchange rate fluctuations.

Equation (4.19.) therefore assumes the form:

$$P_{t+1} - P_t = \frac{R-1}{R-\rho}(\alpha_{Tt+1} - \rho\alpha_{Tt}) - \pi_1(\alpha_{Nt+1} - \alpha_{Nt}) + \pi_2(\alpha_{Nt+1} - \varphi\alpha_{Nt}) \tag{4.22.}$$

Under the assumption that the government's portion of non-traded goods is similar in both open and closed markets (so that ζ_N is identical in both (4.14) and (4.21.) and by assuming that g_N or α_N follow a random walk (φ=1) it may be deduced that shocks to non-traded goods consumption or government expenditure exhibit a similar impact in both open and closed markets. Shocks to productivity in the traded goods sector (α_T), with immovable factors, however exhibit a different effect under closed capital markets on the real exchange rate and these are illustrated as:

$$\mathrm{var}(p_{t+1} - p_t)^{closed} = \sigma_T^2 / (1-\upsilon^2) \tag{4.23.}$$

where σ_T^2 is the var ε_T, and for expediency σ_N^2 is fixed equal to zero. Similarly, the variance of fluctuations in the real exchange rate can be estimated under open capital markets by employing equation (4.21.):

$$\mathrm{var}(p_{t+1} - p_t)^{open} = \left(\frac{R-1}{R-\rho}\right)^2 \sigma_T^2 \tag{4.24.}$$

when ρ=1, the RHS (right-hand side) of (4.23.) as well as of (4.24.) equal σ_T^2. Both are identical when ρ=1 as the conduit of consumption smoothing is inconsequential when productivity shocks are irreversible. It follows that when ρ=0, $\mathrm{var}^{open} < \mathrm{var}^{closed}$, and this instance is consistent for ρ<1. Variation to the real exchange rate is less unpredictable in open capital markets when shocks to productivity are transient since the impact of traded goods consumption shocks are smoothed out so that relative traded and non-traded goods consumption exhibits greater soundness. However, when ρ<1, productivity in the traded goods sector is non-stationary and the capacity to take a loan based on future earnings is amplified and thus too is the impact of shocks to productivity.

Equation (4.18.) can therefore be substituted with the function:

$$\alpha_{Tt} - \alpha_{Tt-1} = (\rho - 1)\alpha_{Tt-1} + \varepsilon_{Tt} \qquad \upsilon<1 \qquad (4.25.)$$

A function which is integrated of order one I(1) or first difference stationary. According to Equation (4.25) variation in the growth rate of productivity is transient however variations to the level are not disposed to smoothing out over the decay of time.

According to Equation (4.25.):

$$\operatorname{var}(p_{t+1} - p_t)^{closed} = \sigma_T^2 / (1 - \upsilon^2) \qquad (4.26.)$$

Unexpected developments in lifetime earnings related to productivity shock must be estimated in order to estimate the variance of real exchange rate adjustments under open capital markets. Equation (4.19) can therefore be replaced by:

$$C_{Tt+1} - C_{Tt} = \frac{R}{R-\upsilon}\left[(a_{Tt+1} - a_{Tt}) - \upsilon(a_{Tt} - a_{Tt-1})\right] \qquad (4.27.)$$

By employing Equation (2.27) to substitute into Equation (4.16.), one produces:

$$\operatorname{var}(p_{t+1} - p_t)^{open} = \left(\frac{R}{R-\upsilon}\right)^2 \sigma_T^2 \qquad (4.28.)$$

By likening the RHS of Equation (4.27) with the RHS of Equation (4.26), we detect that for $0<\upsilon<1$, $\text{var}^{open}<\text{var}^{closed}$. Productivity shocks produce a strong response in traded goods consumption when growth rates are positively correlated so that greater earnings at present are indicative of even greater earnings in the future.

Variance in real exchange rates therefore hinges on the unit processes of the embedded shocks to productivity in the traded goods sector. Transient shocks to productivity in the traded goods sector present the opportunity for agents to exploit open capital markets to smooth traded goods consumption which subsequently prompts smoothing of the relative price of nontraded goods, Equation (4.7.). However, recognition of positive shocks to productivity indicates a higher rate of return on earnings from traded goods productivity present earnings. In this instance, consumption smoothing intensifies variance in the price of non-traded goods.

In conclusion it is discerned that according to the Balassa-Samuelson model where capital and labour are completely interchangeable across sectors that government consumption expenditure exhibits little influence when capital markets are entirely free.

4.2.2. The Standard Real Interest Differential Monetary Model

Monetary models of the exchange rate are founded upon a stable money demand function and the soundness of PPP (Sarno and Taylor, 2002; Taylor, 2007). The real interest differential (RID) model (Frankel, 1979a) relates to monetary fundamentals or assets and assumes complete markets where transaction costs are insignificant. The RID model fashions a relationship between the price of identical assets in two separate economies and the exchange rate. The exchange rate is anticipated to move together with changes in the rates of return on the assets in both economies so that an increased rate of return on a domestic asset prompts a re-valuation of the home currency. Establishing a link between asset prices and the exchange rate is essential to form a comprehensible monetary model.

The RID monetary model assumes uncovered interest rate parity (UIP) and the validity of PPP so that returns on an asset are continuously compounded. The UIP hypothesis assumes that returns on assets, similar in every aspect, equalize across economies so that reward is not required by investors to cover for eventual risk. UIP is represented as:

$$E(\Delta e) = i_t - i_t^* \tag{4.29}$$

Here an * denotes the foreign country, in this model Japan while the USA is home.

While the PPP hypothesis implies that:

$$e_t = p_t - p_t^* \tag{4.30}$$

The RID model can therefore be represented as (for more detail see Chapter 2, section 2.2.1.):

$$e_t = \beta_1(m_t - m_t^*) + \beta_2(y_t - y_t^{s*}) + \beta_3(i_t^s - i_t^{s*}) + \beta_4(i_t^l - i_t^{l*}) + \varepsilon_t \tag{4.31}$$

where e_t is the exchange rate (in this instance USD per unit of JPY), m_t is the money supply, y_t is real income, i_t^s is the short-term interest rate e.g. the discount interest rate and i_t^l is the long-term interest rate e.g. 10-year government bond rate.

The RID model in Eq. (4.31) assumes that the exchange rate is conditioned on money supply at home and abroad so that a rise in the domestic money supply relative to the foreign prompts a rise in domestic prices and a one for one depreciation (β_1=1) of the home currency relative to the foreign currency. Heightened domestic output and decreased prices at home prompt a heightened demand for money and a reduction in domestic prices prompting a re-valuation (β_2<0 and β_4>0) of the domestic currency relative to the foreign currency. An increased domestic interest rate relative to that abroad also triggers a greater supply of money and as a consequence a re-valuation (β_3<0) of the domestic currency relative to the foreign currency. Thus, the model infers that the coefficient on relative money is equal to unity, the coefficients on income as well as the short interest rate are negative and the coefficient on the long-interest rate is positive (Frankel, 1979a). Therefore, when the nominal exchange rate is denominated in domestic terms (USDs) the coefficients are predicted to be β_1=1, β_2<0, β_3<0 and β_4>0.

The standard RID monetary model has historically failed to yield a sound long-run relationship between exchange rates and monetary fundamentals (Hodrick, 1978; Bilson, 1978; Frankel, 1979a; Meese, 1986; Macdonald and Taylor, 1992; Mark, 1995). Hunter and

Ali (2014) however in a hybrid monetary model of the USD/JPY exchange rate manage to establish a long run relationship between the exchange rate and variables related to monetary fundamentals, share prices, traded and non-traded goods productivity and terms of trade. The next section presents an adjusted version of the hybrid monetary model (Hunter and Ali, 2014).

4.2.3. The Adjusted Hybrid Monetary Model

A drawback of the standard RID monetary model is that it accounts for short-term dynamics alone. Friedmann (1988) perceived an inverse long-term relationship between the Dow Jones stock market index and the USA monetary aggregate M2 and consequently concludes that share prices encompass dynamics related to long-run economic expectations. The wealth effect outlines how heightened stock market transactions prompt an increase in the demand for money to hedge greater market risk. The substitution effect however describes how increased share prices result in heightened share returns and a decreased money demand as shares are a substitute for money.

Failure of the standard RID monetary model has in addition been ascribed to the breakdown of PPP. A relationship between the price of traded and non-traded goods and the exchange rate is identified by Harrod (1933) Balassa (1964) and Samuelson (1964) and is dubbed the HBS effect. The HBS effect dictates that increased productivity in the traded goods sector domestically prompts a fall in the price of traded goods as compared to non-traded goods and leads to a re-valuation of the domestic currency. In addition, increased government spending at home triggers a rise in the price of non-traded goods and therefore an appreciation of the domestic currency. Heightened productivity in the traded goods sector is likely to trigger a rise in the cost of labour in both the traded and non-traded goods sectors as labour is interchangeable between sectors. A further refinement of the standard

RID monetary model incorporates a global oil price that accounts for terms of trade and isolates shocks that are exogenous to the system (Rogoff, 1992; Johansen and Juselius, 1992; Kim and Roubini, 2000; Amano and van Nordan, 1998; Lizardo and Mollick, 2010; Hunter and Ali, 2014.)

The hybrid monetary model formulated by Hunter and Ali (2014) is as follows:

$$\begin{aligned} e_t &= \beta_1(m_t - m_t^*) + \beta_2(y_t - y_t^*) + \beta_3(i_t^s - i_t^{s*}) + \beta_4(i_t^l - i_t^{l*}) + \beta_5(s_t - s_t^*) \\ &+ \beta_6(prod_t^T - prod_t^{T*}) + \beta_7(gs_t - gs_t^*) + \beta_8(roil_t^{WTI}) \end{aligned} \tag{4.32.}$$

where e_t is the exchange rate (in this instance domestic currency, dollars, per unit of foreign currency, JPY), m_t is the money supply, y_t is real income, i_t^s is the short-term interest rate e.g. the discount interest rate, i_t^l is the long-term interest rate e.g. 10-year government bond rate, s_t is the stock price index, $prod_t^T$ is productivity in industry, gs_t is real government spending (constant prices) and $roil_t$ is the real price of West Texas Intermediate Crude. Here an * denotes the foreign country, in this model Japan while the USA is home.

For the purpose of this study the hybrid model presented above is adjusted by incorporating the variables in Rogoff's (1992) fixed-factor neoclassical model that examines the relationship between the USD/JPY real exchange rate and traded and non-traded goods productivity. The adjusted hybrid monetary model is as follows

$$\begin{aligned} e_t &= \beta_1(m_t - m_t^*) + \beta_2(y_t - y_t^*) + \beta_3(i_t^s - i_t^{s*}) + \beta_4(i_t^l - i_t^{l*}) + \beta_5(s_t - s_t^*) \\ &+ \beta_6(prod_t^M - prod_t^{M*}) + \beta_7(gs_t - gs_t^*) + \beta_8(roil_t^{WTI}) \end{aligned} \tag{4.33}$$

where $prod_t^M$ is productivity in manufacturing, gs_t is the ratio of government consumption expenditure to GDP at current prices and $roil_t^{Dubai}$ is the real price of Dubai crude.

When the exchange rate is denoted in terms of the domestic currency (here USDs) the coefficients are assumed to be $\beta_1 = 1$ for relative money supply, $\beta_2 < 0$ for relative income, $\beta_3 < 0$ for the relative short-run interest rate, $\beta_4 > 0$ for the relative long-run interest rate, $\beta_5 < 0$, in the case of a wealth effect (capital flight) or $\beta_5 > 0$ for relative real share prices in the case of a substitution effect, $\beta_6 < 0$ for productivity in manufacturing when productivity in the traded goods sectors is greater than that in the non-traded sector, $\beta_7 < 0$ for relative government consumption expenditure when increased government expenditure triggers an increase in the price of non-traded goods and finally would raise the price of nontradables and finally $\beta_8 < 0$ for the real oil price as increased oil prices prompt an appreciation of the U.S dollar. A more detailed description of the hybrid model developed by Hunter and Ali (2014) is presented in Chapter 2 (see section 2.2.2.).

4.3. The Data and Econometric Approach

4.3.1. Data

This analysis involves, for most series, seasonally unadjusted data for the United States vis-à-vis Japan over the period 1980:1 – 2016:4. The study starts in 1980 to take account of financial market deregulation and to probe aspects that determine the USD-JPY exchange rate in the post Bretton Woods period branded by flexible exchange rates. The sample period is characterised by greater capital market integration and financial market upheaval following the mortgage backed securities crisis between 2007 and 2009. We assess the USD/JPY hybrid model constructed by Hunter and Ali (2014) but include the instrumental variables employed by Rogoff (1992). We draw a comparison between the Hunter and Ali (2014) model and the Rogoff (1992) adjusted hybrid model to assess the impact of the global financial crisis of 2008 and the present day that is dominated by zero lower-bound interest rates. Since GDP is not available monthly so we use quarterly data to maintain a clear model.

The exchange rates are the official rates denoted as USD per JPY. The short-term interest rates are represented by the discount rate while the long-term interest rates are 10-year government bond yields. The USA and Japanese consumer price index (CPI) are used to, respectively, deflate the S&P 500 index and the Nikkei 225 Index. Government spending is defined as current government consumption expenditure as a percentage of current gross domestic product while productivity is defined as labour output per hour in manufacturing. The real oil price is the spot price of Dubai Crude, dollars per barrel, deflated by the USA consumer price index. The exchange rates, interest rates, wholesale and consumer price indices (CPI) are extracted from the IMF's International Financial Statistics (IFS). Money supply (M1) data as well as stock and oil prices are collected from Thomson DataStream. Both nominal and real gross domestic product and final consumption figures for the USA as well as Japan are obtained from the OECD main economic indicators (MEI) database. Labour productivity data for Japan is retrieved from the Japan Productivity Center[19] while corresponding data for the USA is obtained from the USA Bureau of Labor Statistics.

To calculate the government spending ratio, we employ gross domestic product and final government consumption expenditure figures that are expenditure based, at current prices and volume estimates chained to the OECD base year 2010. Exchange rates, interest rates, stock prices and the oil price are end of period data. All variables are expressed in their logarithmic form except interest rate differentials and the government spending ratio that are a percentage. Finally, the graphs of the variables for the structural (Rogoff, 1992), standard RID (Frankel, 1979a) and adjusted hybrid monetary models of the USD/JPY

[19] http://www.jpc-net.jp/eng/stats/

exchange rate in levels (upper graphs) and first differences (lower graphs) are displayed in Figures 4.8.-4.18. below.

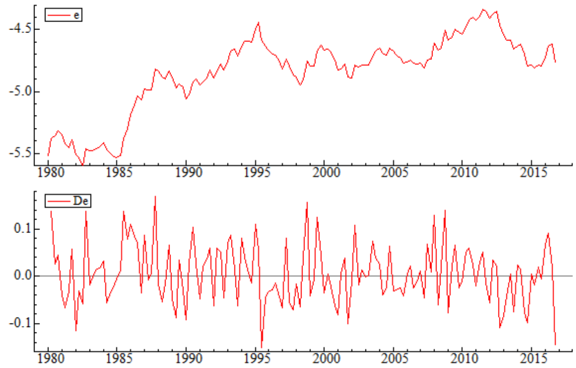

Figure 4.8. Spot USD/JPY exchange rate

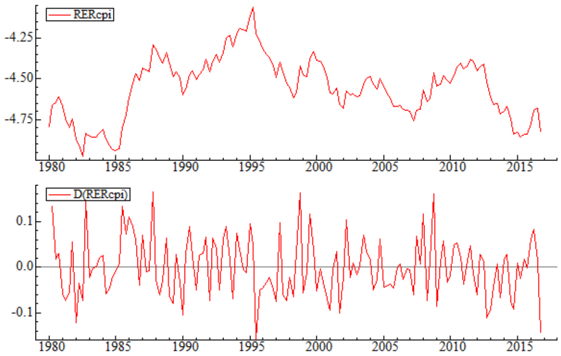

Figure 4.9. The real spot USD/JPY^{CPI} exchange rate

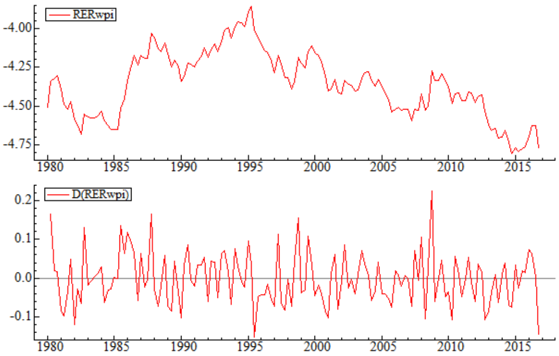

Figure 4.10. The real spot USD/JPY^{WPI} exchange rate

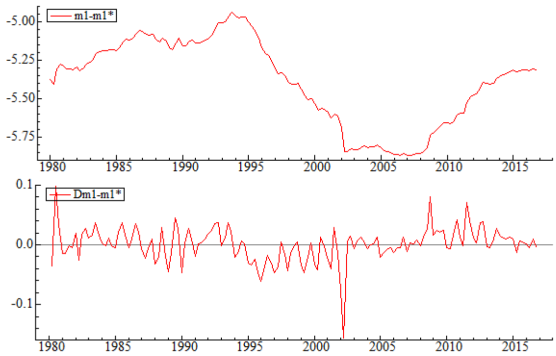

Figure 4.11. Relative money supply (M1-M1*)

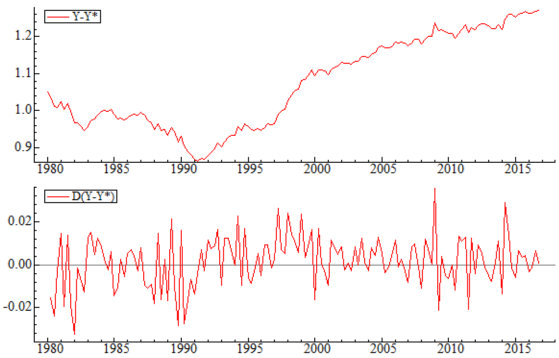

Figure 4.12. Relative income (Y-Y*)

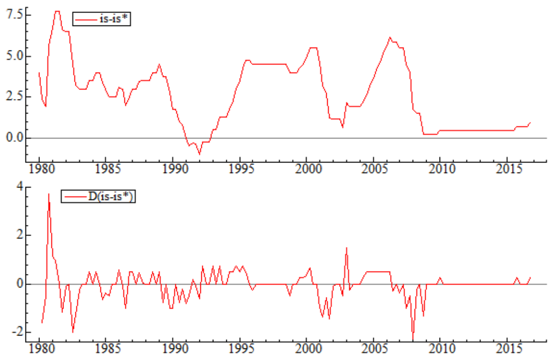

Figure 4.13. Short-term interest differential

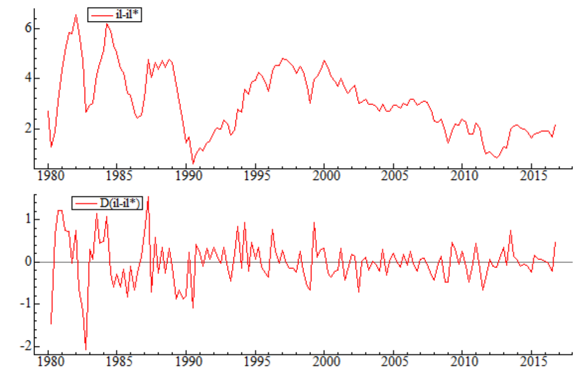

Figure 4.14. Long-term interest differential

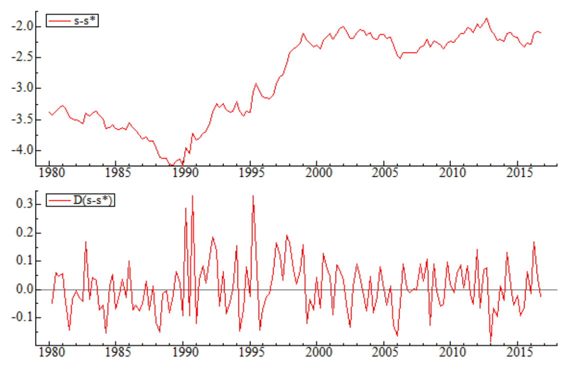

Figure 4.15. Relative real stock prices

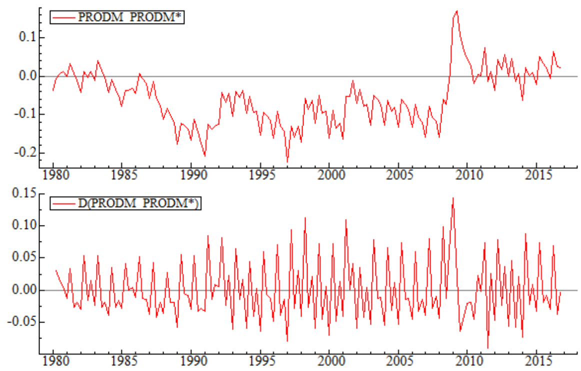

Figure 4.16. Productivity differential

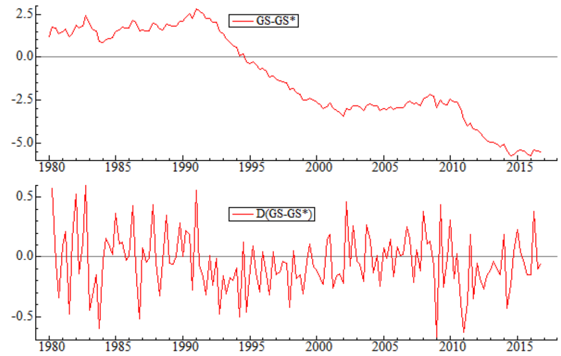

Figure 4.17. Government spending differential

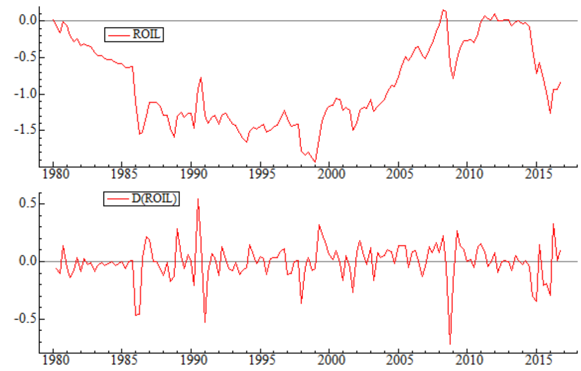

Figure 4.18. Real oil price

4.3.2. Tests of Stationarity

A condition for performing cointegration tests is that the series under scrutiny are integrated of order one I(1). By means of an Augmented Dicky-Fuller (ADF) (Dickey and Fuller, 1981) test we show that all the variables in our sample are non-stationary in levels or I(0) and first difference stationary or I(1)). The ADF test results for the variables employed in Rogoff's (1992) structural model and the variables used in Frankel's (1979a) traditional RID model of the USD/JPY exchange rate are shown in Table 4.1. and Table 4.2., respectively below. Lastly, ADF unit root test results for the variables employed in the adjusted hybrid model of the USD/JPY exchange rate are shown in Table 4.3., also below. The ADF test results confirm that the variables in our sample are first difference stationary for a 95% significance level. We allow for a maximum of four lags as our data is quarterly.

Table 4.1. ADF Unit Root Test Results for Rogoff's (1992) Structural Model of the USD/JPY Exchange Rate

$$\Delta^2 x_t = \pi_0 + \gamma \Delta x_{t-1} + \pi_1 \Delta^2 x_{t-1} + \varepsilon_t$$

Variable	Level		First differences
	Constant	Constant and trend	Constant
real exchange rate^{CPI}	-1.99(0)	-1.95(0)	-4.68(3) **
real exchange rate^{WPI}	-1.45(0)	-2.00(0)	-5.09(3) **
prod^{M}-prod^{M*}	-1.82(3)	-2.11(3)	-4.49(3) **
gs-gs*	0.29(0)	-2.20(0)	-5.76(2) **
roil	-1.80(2)	-2.15(2)	-10.19(0) **

Note: the 1% and 5% critical values for the ADF tests are respectively -3.49 and -2.89 and (without trend) and -4.02 and -3.44 (with the trend); the proper lag length, allowing for a maximum of four lags is selected by the modified Schwarz information criterion; ** and * indicate statistical significance at the 1% and 5% levels, respectively.

Table 4.2. ADF Unit Root Test Results for Frankel's (1979a) Traditional RID Model of the USD/JPY Exchange Rate

$$\Delta^2 x_t = \pi_0 + \gamma \Delta x_{t-1} + \pi_1 \Delta^2 x_{t-1} + \varepsilon_t$$

Variable	Level		First differences
	Constant	Constant and trend	Constant
e	-2.19(0)	-1.81(0)	-4.62(3) **
$m - m^{*}$	-1.64(4)	-1.66(4)	-2.98(3)*
$y - y^{*}$	-0.06(3)	-2.76(0)	-5.12(2) **
$i^{s} - i^{s*}$	-1.85(0)	-2.08(0)	-5.61(2) **
$i^{l} - i^{l*}$	-2.33(0)	-3.02(2)	-6.67(1) **

Note: the 1% and 5% critical values for the ADF tests are respectively -3.49 and -2.89 and (without trend) and -4.02 and -3.44 (with the trend); the proper lag length, allowing for a maximum of four lags is selected by the modified Schwarz information criterion; ** and * indicate statistical significance at the 1% and 5% levels, respectively.

Table 4.3. ADF Unit Root Test Results for the Adjusted Hybrid Model of the USD/JPY Exchange Rate			
$\Delta^2 x_t = \pi_0 + \gamma \Delta x_{t-1} + \pi_1 \Delta^2 x_{t-1} + \varepsilon_t$			
Variable	Level		First differences
	Constant	Constant and trend	Constant
e	-2.19(0)	-1.81(0)	-6.040501(2) **
real exchange rate^{CPI}	-1.99(0)	-1.95(0)	-4.68(3) **
real exchange rate^{WPI}	-1.45(0)	-2.00(0)	-5.09(3) **
$m - m^*$	-2.19(0)	-1.81(0)	-4.62(3) **
$y - y^*$	-1.64(4)	-1.66(4)	-2.98(3)*
$i^s - i^{s*}$	-0.06(3)	-2.76(0)	-5.12(2) **
$i^l - i^{l*}$	-1.85(0)	-2.08(0)	-5.61(2) **
$s\text{-}s^*$	-0.63 (0)	-1.72(0)	-7.59(1) **
$prod^M\text{-}prod^{M*}$	-1.82(3)	-2.11(3)	-4.49(3) **
$gs\text{-}gs^*$	0.29(0)	-2.20(0)	-5.76(2) **
roil	-1.80(2)	-2.15(2)	-10.19(0) **
Note: the 1% and 5% critical values for the ADF tests are respectively -3.49 and -2.89 and (without trend) and -4.02 and -3.44 (with the trend); the proper lag length, allowing for a maximum of four lags is selected by the modified Schwarz information criterion; ** and * indicate statistical significance at the 1% and 5% levels, respectively.			

4.3.3. The Econometric Approach

The Johansen (1995) methodology is a development of the Engle and Granger (1987) two step methodology for identifying long-run relationships. Johansen (1988; 1995) introduced the Maximum Likelihood test as a more robust approach to establishing equilibrium relationships amongst variables that are level non-stationary in a multivariate setting. The focus of the methodology is to recognize stochastic trends and cointegrating relationships and examine their composition. Hoover et al. (2008) advocate the approach firstly, because non-stationary economic series are known to be problematic and secondly because they focus on the adaptation of one variable to another and the search for the most favourable outcome for a single factor as well as a congruent structural effect. The Johansen (1995) methodology has been employed in a vast number of studies and is endorsed with regard to monetary models of the exchange rate (see Macdonald, 2007, Chapter 6). Widely recognized are studies conducted by Macdonald and Taylor (1993 and

1994), Sarantis (1994), McNown and Wallace (1994), Chinn (1997a) and Cheung and Chinn (1998). Johansen (1995) assumes an unrestricted vector autoregression model of order p with n I(1) endogenous variables that is driven by a vector X_t with (n x 1) Gaussian errors expressed as an error correction as follows:

$$\Delta X_t = \Pi X_{t-1} + \Gamma_1 \Delta X_{t-1} + ...+ \Gamma_p \Delta X_{t-(p-1)} + \gamma D_1 + \varepsilon_1, \tag{4.34}$$

where X_t is an (n × 1) vector of variables tied to the theoretical long-run link originating from monetary models, D_1 is a vector of constants, seasonal dummies, and impulse dummies; Γ_i (i =1 ,....., p-1) are (n × n) variable matrices encapsulating the short-run dynamics between model variables, and **Π** is an (n × n) matrix that is structured as **Π**=**αβ**′, where **α** denotes the rate of adjustment and **β** matrices denote elements with a long-run trend.

The trace test and maximum eigenvalue statistic are constituents of the maximum likelihood tests introduced by Johansen (1995) to examine cointegration and establish the rank order r of the long-run matrix **Π**. For the purpose of this analysis the trace test is implemented to establish the rank order r of Π. The Johansen (1995) test follows a superlative succession and begins with the null hypothesis of zero cointegrating vectors r=0 against the alternative of at least a single cointegrating vector r ≤1 and progresses with the elimination of supplementary ranks of cointegration until r = i against the alternative r ≤ i+1; the succession halts at r = i as the null can no longer be rejected. Hubrich et al. (2001) find that imposing restrictions by cointegrating rank is ineffectual for forecasts at short intervals however Engle and Yoo (1987) and Clements and Hendry (1998) find that employing an unrestricted VAR at long-horizons yields powerful results. The trace test as expressed in terms of eigenvalues (λ_i) and sample size (T) is represented as follows:

$$\lambda_{trace} = -T \sum_{i=1}^{n} (1-\lambda_i). \tag{4.35}$$

The success of the Johansen (1995) test lies with the sound establishment of a well formulated VAR as the Johansen (1995) test for cointegration is founded on an unrestricted VAR (Johansen, 1995). Lag length is most frequently determined by information criteria namely the Bayesian information criterion, Schwarz criterion or Akaike information criterion. Burke and Hunter (2005) argue that inadequate choice of lag length can misrepresent the size of the trace test with respect to the null distribution so suggest amplifying the number of lags when serial correlation is detected.

We employ the procedure described above to distinguish a persistent long-run equilibrium relation for the USD/JPY exchange rate. Eq. (4.34) is used to estimate the standard RID monetary model and an adjusted hybrid monetary model, expressed as the vectors in the corresponding rows:

$$X'_{(RID)t} = [e_t, m_t - m_t^*, y_t - y_t^*, i_t^s - i_t^{s*}, i_t^l - i_t^{l*}],$$

$$X'_{(ADJUSTED\_HYBRID)} = [e_t, m - m_t^*, y_t - y_t^*, i_t^s - i_t^{s*}, i_t^l - i_t^{l*}, s_t - s_t^*, prod_t^M - prod_t^{M*}, gs_t - gs_t^*, roil_t]$$

The aim of this analysis is to identify elements with a deterministic trend and to substantiate the long-run relation implicit in the model. First root properties of the individual series are investigated, and the series are expected to be non-stationary in levels. The Johansen (1995) procedure is subsequently employed in order to identify a meaningful long-run relationship. Tests of LE (Juselius, 1995) and WE (Johansen, 1992b) are applied to ensure soundness of the model as a long-run explanation of the exchange rate. Variables that are found to be weakly exogenous are perceived to propel the system rather than adjust to it in the long-run. Burke and Hunter (2005) claim that recognition of long-run excluded and weakly exogenous variables not only aids in identification of a long run relation but provides insight into the dynamics of such a relation.

4.3.4. The Heterogeneous Case for Unit-Root Testing

To test for stationarity we apply the Augmented Dickey-Fuller test for stationarity to each variable in our sample allowing for a maximum of four lags considering that the data is quarterly. The ADF test assumes the following representation:

$$\Delta\bar{x}_{it} = \mu_i + \psi_i \bar{x}_{it-1} + \sum_{j=1}^{p-1} \psi_j \Delta\bar{x}_{it-j} + \varepsilon_{it} \qquad (4.36)$$

where $\bar{x}_{\mathrm{it}} = x_{\mathrm{it}} - \sum_{\mathrm{j=1}}^{\mathrm{t}} \frac{x_{ij}}{t}$

The large sample distribution of the ADF test is not sensitive to the inclusion of differenced time series and controls for the initial values suggesting that each series is homogeneous relative to the point of departure. Eq (4.36) is estimated by OLS and critical values for comparison of Eq. (4.36) are calculated under the null ($\psi_i=0$). The variance is corrected to account for more heterogeneity as well as the model estimated by maximum likelihood to compute the autoregressive conditional heteroskedasticity (ARCH)[20].

4.4. Empirical Results

4.4.1. Results for Rogoff's 1992 Structural Model

This analysis commences by considering the structural model proposed by Rogoff (1992) under the assumption that the exchange rate follows a random walk and that other components of the model are shocks. Meese and Rogoff (1983a) however evidence that this type of model does not beat the out-of-sample naive random walk and cannot

[20] Beirne et al. (2007) show by application of an ADF test where the t-statistics employ White standard errors for all but two countries in their sample, where the ARCH method is employed, that eight out of the ten countries in the sample exhibit stationary real exchange rates at a nominal 5% significance level. The standard errors for Luxemburg and Portugal are extracted from a model that estimates ARCH(1) and ARCH(4) terms in the variance. The ARCH(1) and ARCH(4) estimates are considered more efficient and thus the test statistics more exact on condition that the volatility is consistently well behaved. Considering this the traditional large sample or asymptotic critical values may be adopted.

explain dynamics that drive the exchange rate. Rogoff's (1992) structural model is investigated as the status quo has changed since its development and a systems equation methodology does not consider endogeneity since information is assumed to be predetermined. This analysis considers the nature of cointegrating relationships in the current age of flexible exchange rate regimes, the duration of the worst financial crisis since the Great Depression, 2007-2009, and the subsequent period of zero lower bound interest that has been extended for almost a decade. Instrumental variables are employed in Rogoff's (1992) structural model to establish a long-run exchange rate relation. A necessary condition for carrying out cointegration tests is that the series under investigation are integrated of order one I(1). A finding of cointegration can assist in understanding and identifying a long-run relation (Burke and Hunter, 2005, ch5) that depicts an exchange rate equation. Likelihood ratio tests should demonstrate that the exchange rate is neither long-run excluded (Juselius, 1995), nor weakly exogenous (Johansen, 1992b). Augmented Dicky-Fuller (ADF) (Dickey and Fuller, 1981) unit root tests are applied to examine the time series properties of variables included in the structural model (Rogoff, 1992). ADF test results for the sample series are presented in Tables 4.1. - 4.3. in section 4.3.2. Unit root test results confirm that all sample variables are first difference stationary and therefore I(1).

Cointegration tests and tests of LE, WE and stationarity (S) for models of the real USD/JPY^{CPI} exchange rate including a single explanatory variable are exhibited in Tables 4.4. – 4.9., below. Table 4.4. illustrates that the real exchange rate, deflated by CPI, and the measure for government consumption expenditure are co-integrated however, Table 4.5. (see Panel B) shows that government consumption expenditure can be long-run excluded from this model of the real exchange rate. Table 4.6. illustrates that the real exchange rate, deflated by CPI, and productivity in manufacturing are cointegrated while Table 4.7. (see

Panel B) illustrates that the real CPI exchange rate is weakly exogenous to this system. Table 4.8. demonstrates that at least 2 cointegrating relations exist, for a 90% significance level, in the model comprising the real USD/JPY^{CPI} exchange rate and the real oil price while Table 4.9. (see Panel B) shows that the real oil price is exogenous to the system.

Table 4.4. Test that the Real USD/JPY^{CPI} Exchange Rate and Government Consumption Expenditure are Cointegrated. The System Comprises of [*RER^{CPI} and gs-gs]**						
(p –r)	*rank*	Eigenvalue	Trace Test	90% Critical Value	95% Critical Value	*p-values*
2	*r = 0*	0.09	16.5	13.33	15.41	[0.03] **
1	*r ≤ 1*	0.02	2.19	2.69	3.84	[0.14]
Notes: *r* denotes the number of cointegrating vectors. ** and * indicates statistical significance at the 5% level and 10% level, respectively. p-values in brackets [ ].						

Table 4.5. LE, WE, and Stationarity (S) Tests for the USD/JPY^{CPI} and Government Consumption Expenditure Ratio Model		
	USD/JPY^{CPI}	gs-gs*
Panel A: LE tests		
$X^2(1)$	11.09	2.07
p-values	[0.00]**	[0.15]
Panel B: WE tests		
$X^2(1)$	6.6289	8.16
p-values	[0.01]*	[0.00]**
Panel C: S tests		
$X^2(1)$	2.07	11.09
p-values	[0.15]	[0.00]**
Notes: ** and * indicate significance at the 1% and 5% levels, respectively. p-valuess are in brackets [].		

Table 4.6. Test that the real USD/JPY^{CPI} and Productivity in Manufacturing are Cointegrated. The System Comprises of [*RER^{CPI} and prod^{M}-prod^{M*}*]						
(p –r)	*rank*	Eigenvalue	Trace Test	90% Critical Value	95% Critical Value	*p-values*
2	*r = 0*	0.09	17.69	13.33	15.41	[0.02] **
1	*r ≤ 1*	0.015	2.28	2.69	3.84	[0.13]
Notes: r denotes the number of cointegrating vectors. ** and * indicates statistical significance at the 5% level and 10% level, respectively. p-values in brackets [ ].						

Table 4.7. LE, WE, and Stationarity (S) Tests for the USD/JPY^{CPI} and Productivity in Manufacturing Model		
	USD/JPY^{CPI}	$prod^{M}$-$prod^{M*}$
Panel A: LE tests		
$X^2(1)$	8.32	9.00
p-value	[0.00]**	[0.00]**
Panel B: WE tests		
$X^2(1)$	2.16	10.23
p-value	[0.14]	[0.00]**
Panel C: S tests		
$X^2(1)$	11.091	7.7100
p-value	[0.00]**	[0.01]**
Notes: ** and * indicate significance at the 1% and 5% levels, respectively. p-values are in brackets [].		

Table 4.8. Test that the Real USD/JPY^{CPI} and the Real Oil Price are Cointegrated. The System Comprises of [*RER^{CPI} and roil*]					
(p –r)	*rank*	Eigenvalue	Trace Test	90% Critical Value	*p-value*
2	*r = 0*	0.09	13.61	13.3	[0.09] *
1	*r ≤ 1*	0.02	3.93	2.69	[0.05] *
Notes: r denotes the number of cointegrating vectors. ** and * indicates statistical significance at the 5% level and 10% level, respectively. p-values in brackets [ ].					

Table 4.9. LE, WE, and Stationarity (S) Tests for the USD/JPY^{CPI} and the Real Oil Price Model Where a Single Cointegrating Vector is Identified		
	USD/JPY^{CPI}	roil
Panel A: LE tests		
$X^2(1)$	4.15	4.62
p-value	[0.04]*	[0.03]*
Panel B: WE tests		
$X^2(1)$	5.73	0.09
p-value	[0.02]*	[0.763
Panel C: S tests		
$X^2(1)$	4.62	4.15
p-value	[0.03]*	[0.04]*
Notes: ** and * indicate significance at the 1% and 5% levels, respectively. p-values are in brackets [].		

Since the models depicted above are not feasible a further model of the real CPI exchange rate is formulated to include all the variables related to traded and non-traded goods productivity namely, productivity in manufacturing, government consumption expenditure and the real oil price. A test of cointegration for the comprehensive traded and non-traded goods productivity model of the real USD/JPY^{CPI} exchange rate is presented below in Table 4.10. (Misspecification tests for the model are presented in Appendix A4.2. Table A4.1.). The cointegration test suggests that, for a 95% significance level, a single long-run cointegrating relationship exists between the real USD/JPY^{CPI} exchange rate and the explanatory variables for government consumption expenditure, productivity in manufacturing and the real oil price.

Table 4.10. Test That the Real USD/JPY^{CPI} and Variables Related to Traded and Non-Traded Goods Productivity are Cointegrated. The System Comprises of [RER^{CPI}, $prod^{M}$-$prod^{M*}$, gs-$gs*$ and $roil$]

(p –r)	*rank*	Eigenvalue	Trace Statistic	90% Critical Value	95% Critical Value	*p-value*
4	$r = 0$	0.20	51.82	43.95	47.21	[0.00] **
3	$r \leq 1$	0.2	24.29	26.79	29.68	[0.52]
2	$r \leq 2$	0.81	4.97	13.33	15.41	[0.75]
1	$r \leq 3$	0.26	1.25	2.69	3.76	[0.81]

Notes: r denotes the number of cointegrating vectors. ** and * indicate statistical significance at the 5% and 10% level, respectively. p-values in brackets [ ].

The estimated cointegrating vector for the comprehensive traded and non-traded goods productivity model of the real USD/JPY^{CPI} exchange rate is presented below in Table 4.11. and illustrates that the coefficient on relative productivity in manufacturing is significant at the 1% level whilst the others are insignificant. Tests of LE and WE presented in Table 4.12. below show that the real USD/JPY^{CPI} exchange rate and the real oil price are weakly exogenous to the system but that the real oil price is exogenous both in terms of WE and strict exogeneity. The term for government consumption expenditure can in addition to the real oil price be long-run excluded from the framework. WE in terms of the real USD/JPY^{CPI} exchange rate suggests that it does not adapt to a long-run equilibrium but drives the framework instead. This finding implies that the comprehensive traded and non-traded goods productivity model is incoherent and that the real USD/JPY^{CPI} exchange rate follows a random walk. Joint tests of LE and WE for roil and gs-gs* can be found in Appendix A4.2. (see Tables A4.2. and A4.3., respectively).

Table 4.11. The Estimated Cointegrating Vector for the Traded and Non-Traded Goods Productivity Model Normalised on the Real USD/JPY^{CPI} Exchange Rate			
	prod^{M}-prod^{M*}	gs-gs*	roil
Coefficient	4.56	0.02	-0.08
Standard error	0.84	0.02	0.09
t-stat	5.43**	1.4	-0.83
Notes: ** and * indicate significance at the 1% and 5% levels, respectively based on standard inference. p-values are in square brackets[.].			

Table 4.12. LE, WE, and Stationarity (S) Tests for the Traded and Non-Traded Goods Productivity Model of the Real USD/JPY^{CPI} Exchange Rate				
What	RER	prod^{M}-prod^{M*}	gs-gs*	roil
Panel A: LE tests				
$X^2(1)$	11.02	22.33	2.26	0.54
p-value	[0.00]**	[0.00]**	[0.13]	[0.46]
Panel B: WE tests				
$X^2(1)$	0.98	17.86	5.40	0.85
p-value	[0.32]	[0.00]**	[0.02]*	[0.36]
Panel C: S tests				
$X^2(1)$	31.74	22.33	40.8	36.03
p-value	[0.00]**	[0.00]**	[0.00]**	[0.00]**
Notes: ** and * indicate significance at the 1% and 5% levels, respectively. p-values are in brackets [].				

For robustness we estimate the comprehensive traded and non-traded goods productivity model of the real USD/JPY exchange rate deflated by WPI as compared to CPI. Tests of cointegration in Table 4.13., below, illustrate that a single cointegrating exists amongst model variables and the real USD/JPY^{WPI} exchange rate. However, tests of LE in Table 4.14., Panel A, confirm that this is not a model of the exchange rate as the variable of primary interest, the real exchange rate, is not different from zero, it is weakly exogenous and can be long-run excluded. In addition, tests of WE in Table 4.14., Panel B, show that the real USD/JPY^{WPI} exchange rate is exogenous in terms of both weak and strict exogeneity.

The tests of exogeneity imply that the real USD/JPY^{WPI} exchange rate drives the long-run equilibrium as rather than adapting to it. Therefore, a long-run relation based on this model of the USD/JPY^{WPI} exchange rate would be incongruent. Misspecification tests, the estimated cointegrating relation as well as joint tests of LE and WE for roil and gs-gs* for the USD/JPY^{WPI} model are presented in Appendix A4.2. (See Tables A4.4., A4.5., A4.6. and A4.7., respectively.).

Table 4.13. Test That the Real USD/JPY^{WPI} and Variables Related to Traded and Non-Traded Goods Productivity are Cointegrated. The System Comprises of [*RER^{CPI}, prod^{M}-prod^{M*}, gs-gs and *roil*]**

(p –r)	*rank*	Eigenvalue	Trace Statistic	95% Critical Value	*p-value*
4	*r = 0*	0.18	54.34	47.21	[0.01] **
3	*r ≤ 1*	0.13	24.90	29.68	[0.17]
2	*r ≤ 2*	0.03	4.71	15.41	[0.84]
1	*r ≤ 3*	0.00	0.05	3.76	[0.82]

Notes: *r* denotes the number of cointegrating vectors. ** Indicates statistical significance at the 1% level. P-values in brackets [ ].

Table 4.14. LE, WE, and Stationarity (S) Tests for the Traded and Non-Traded Goods Productivity Model of the USD/JPY^{WPI}

	RER	prod^{M}-prod^{M*}	gs-gs*	roil
Panel A: LE tests				
$X^2(1)$	3.58	7.23	0.16	0.01
p-value	[0.06]	[0.01]**	[0.69]	[0.93]
Panel B: WE tests				
$X^2(1)$	0.70	7.31	2.34	0.02
p-value	[0.40]	[0.01]**	[0.13]	[0.90]
Panel C: S tests				
$X^2(1)$	24.21	17.00	26.88	24.67
p-value	[0.00]**	[0.00]**	[0.00]**	[0.00]**

Notes: ** and * indicate significance at the 1% and 5% levels, respectively. p-values are in brackets [].

The nature of these findings further illustrates the possibility that the exchange rate may follow a random walk and as such drives the system in the long run. It is to be expected from the definition of exogeneity (Ericsson and Irons, 1994) that tests of WE depend on the model in which they are embedded. Therefore, a model that captures a significant portion of fundamental drivers has the potential to reverse the finding that the exchange rate follows a random walk.

In the following section the USD/JPY exchange rate is modelled in the context of international finance by employing variables that relate to stable money demand and the validity of PPP in the traditional RID monetary model (Frankel, 1979a).

4.4.2. Results of the RID Monetary Model

The analysis continues with an empirical test of the classical RID monetary model (Frankel, 1979a) of the USD/JPY exchange rate. The VAR model comprises the $X'_{(RID)t}$ vectors and the variables included in this model are found to be first difference stationary. ADF unit roots tests for the variables used in the RID model are presented in Table 4.2. in Section 4.3.2. Proper lag length is chosen by estimation of a preliminary VAR of 4 lags and reference to the significance of F-tests on retained regressors. The F-tests imply that the correct number of lags for the RID model of the nominal USD/JPY exchange rate is four. The VAR specification for the conventional RID model of the USD/JPY exchange rate comprises a constant, three seasonal dummy variables and impulse dummy variables pertaining to the USA saving and loan crisis in the earlier 1980's (1982Q1), the dot.com stock market crash in mid-2002 (2002Q2) and the pinnacle of the mortgage backed securities crisis in 2008 (2008Q1 and 2008Q4). The implied misspecification tests of the VAR for the RID model of the USD/JPY exchange rate at lag 4 are reported in Table A4.8. in Appendix A4.2.

Table A4.8. shows that, at the 1% significance level, the USD/JPY RID model is free from serial correlation, by means of LM tests of order (8), and ARCH effects, by means of ARCH tests of order (8). The normality tests show that aside from the discount rate differential, all variable residuals are normally distributed at a significance level of 1%. The vector of residuals from the VAR for the USD/JPY exchange rate is in addition normal and free from serial correlation at the 5% significance level. The USD/JPY RID model appears to be well specified despite evidence of non-normality for the discount rate differential at a 1% significance level.

Table 4.15. Johansen Cointegration Test Results for the USD/JPY RID Model The system comprises of [e, m – m*, y – y*, $i^s - i^{s*}$, $i^l - i^{l*}$]						
(p –r)	*rank*	Eigenvalue	Trace Test	95% Critical Value	99% Critical Value	*p-value*
5	*r = 0*	0.30	97.37	68.52	76.07	[0.00] **
4	*r ≤ 1*	0.19	46.07	47.21	54.46	[0.07]
3	*r ≤ 2*	0.08	16.60	29.68	35.65	[0.68]
2	*r ≤ 3*	0.03	5.25	15.41	20.04	[0.78]
1	*r ≤ 4*	0.01	1.46	3.76	6.65	[0.23]
Notes: *r* denotes the number of cointegrating vectors. * and ** indicate statistical significance at the 1% and 5% level, respectively. P-values in brackets [ ].						

The estimated eigenvalues and trace statistics for the standard RID monetary model of the USD/JPY exchange rate are illustrated in Table 4.15., a single cointegrating vector is distinguished among the model variables so that the null hypothesis of no cointegration is rejected for the conventional RID monetary model of the USD/JPY exchange rate. The identified cointegrating relation, where rank(Π)=1, is estimated by normalising on the exchange rate which is the variable of primary interest and is therefore expected to be endogenous to the system. Boswijk (1996) however, advocates that an additional rank condition be applied to examine the soundness of identifying restrictions. Normalising on a stationary variable is not viable since a cointegrating vector defines a linear combination of nonstationary series (Burke and Hunter, 2005, ch5). The estimated cointegrating vector for the RID model of the USD/JPY exchange rate is presented in Table 4.16. as follows:

Table 4.16. The Estimated Cointegrating Vector for the RID Model of the USD/JPY Exchange Rate				
	m-m*	y-y*	$i^s - i^{s*}$	$i^l - i^{l*}$
Coefficient	0.42	0.29	0.1	0.22
Standard error	0.23	0.54	0.04	0.06
t-stat	1.81*	0.53	2.63**	3.74**
Notes: ** and * indicate significance at the 1% and 5% levels, respectively based on standard inference. p-values are in square brackets[.].				

To reiterate, the nominal USD/JPY exchange rate in this RID model is denoted in terms of the domestic currency (USDs) so that a coefficient rise for the domestic currency would indicate a depreciation of the USD as compared to an appreciation. It is noted that the coefficient on the money supply differential is positive (β_1>0) and significant at the 1% level thus a surge in the relative money supply leads to a depreciation of the domestic currency in this model. The outcome opposes the theory that a rise in domestic money supply relative to the foreign counterpart heightens domestic prices and prompts a one-for-one depreciation of the domestic currency since the derived coefficient is different to unity. The coefficient on the income differential is insignificant and positive (β_2>0). A positive coefficient rejects the theory that a rise in domestic income relative to foreign income raises the demand for the domestic currency and brings about an appreciation of the domestic currency. The coefficient on the discount rate differential is positive (β_3>0) and significant at the 5% level and rejects the theory that a surge in capital mobility toward the domestic economy relative to the foreign prompts an appreciation of the domestic currency. The coefficient on the long-run interest rate differential is positive (β_4>0), at the 5% significance level, and is evident of a substitution effect where the return on stocks is higher than the return on money so investors substitute money for shares and this decreases the demand for domestic currency relative to foreign and triggers a depreciation of the domestic currency. All variable coefficients are of the wrong sign while the coefficient on relative money is far from unity and that on relative income is insignificant. This result echoes that derived by Hunter and Ali (2014) who cannot empirically confirm that the RID model is a model of the exchange rate.

Results of asymptotically chi-squared (Johansen, 1992b; Asteriou and Hall, 2015), tests of LE, WE, and stationarity for the USD/JPY RID model are presented in Table 4.17., below. The tests of WE and LE show that the exchange rate operates like a random walk so cannot predict fundamentals that drive the framework (Engel and West, 2005) in the RID model.

Table 4.17. LE, WE, and Stationarity (S) Tests for the Standard RID Model of the USD/JPY Exchange Rate					
	e	m-m*	y-y*	$i^s - i^{s*}$	I^l-i^{l*}
Panel A: LE tests					
$X^2(1)$	19.51	2.77	0.3	3.24	5.12
p-value	[0.00]**	[0.1]	[0.58]	[0.07]	[0.02]*
Panel B: WE tests					
$X^2(1)$	3.41	14.73	0.09	6.22	3.62
p-value	[0.06]	[0.00]**	[0.76]	[0.01]*	[0.06]
Panel C: S tests					
$X^2(1)$	40.30	45.61	44.17	25.22	25.27
p-value	[0.00]**	[0.00]**	[0.00]**	[0.00]**	[0.00]**
Notes: ** and * indicate significance at the 1% and 5% levels, respectively. p-values are in brackets [].					

The LE tests in Table 4.17., Panel A, suggest that the relative money supply, relative income and relative discount rate can be excluded from the cointegrating system. Tests of WE in Table 4.17., Panel B, suggest that the exchange rate, relative income and relative long-run interest rate are weakly exogenous. A model therefore based on the exchange rate cannot be interpreted since normalisation on an exogenous factor within the cointegrating framework is inapplicable. Hence, the RID model as a long-run model of the USD/JPY exchange rate is incompatible. The exchange rate in this framework behaves like a random walk so does not adapt to the long run equilibrium but moves in response to systemic shocks instead (Hunter, 1992). This conclusion is analogous to that of Hunter and Ali (2014) for the RID model of the nominal USD/JPY exchange rate for the sample period 1980Q1-2009Q4. The term structure of interest rates is upheld in this framework since WE is negated for the short-term interest rate differential but not for the long-term interest rate differential so the course of transmission stems from the short-term interest rate to the long-term interest rate. Juselius and Macdonald (2004) illustrate an inverse phenomenon in a parity model for the USD/JPY exchange rate and suggest that the long-term interest rate is directed by the short-term interest rate. We propose enhancing the system with variables related to non-traded and traded goods productivity and share prices to reverse causality in the traditional RID monetary model of the USD/JPY exchange rate. Stationarity tests in the multivariate framework (see Table 4.17., Panel C) affirm that all the variables in the system are level nonstationary. Joint tests of WE and LE for the differentials relating to income, money and the short-term interest rate can be found in Appendix A4.2. (Tables A4.9-A4.11).

Upon contemplation of the monetary hypothesis it is concluded that the RID monetary model of the USD/JPY exchange rate, for the period 1980Q1-2016Q4, is an incompatible framework for modelling the long-run equilibrium exchange rate since explanatory variables cannot be normalised on the exchange rate as it is shown to be exogenous to the cointegrating system. The nominal USD/JPY exchange rate in this instance appears to behave like a random walk.

4.4.3. Results of the Adjusted Hybrid Monetary Model

The final part of this analysis further tests the long-run relations established earlier in the traded goods consumption smoothing structural model (Rogoff, 1992) and the standard RID monetary model (Frankel, 1979a). An adjusted hybrid monetary model, akin to the hybrid model developed by Hunter and Ali (2014), is constructed by incorporating the instrumental variables employed in the traded goods consumption smoothing model of the real USD/JPY exchange rate (Rogoff, 1992). The adjusted hybrid model also includes variables relating to money (Frankel, 1979a), real stock prices (Friedmann, 1988) and a real global oil price that accounts for exogenous shocks related to terms of trade (Johansen and Juselius, 1992; Amano and Norden, 1998). The proposed model diverges from the hybrid model developed by Hunter and Ali (2014) by considering traded goods productivity in manufacturing as compared to in industry and by employing the real price of Dubai Crude as compared to the real price of WTI Crude since most of Japan's oil imports originate from the Middle East[21].

The standard RID model of the USD/JPY exchange rate demonstrated inefficiency as a long-run explanation of the nominal USD/JPY exchange rate, likely a result of unstable money demand and the collapse of PPP. The adjusted hybrid model of the USD/JPY considers stable money demand functions and the soundness of PPP by considering cross-country productivity differentials that affect the price of traded and non-traded goods (Balassa, 1964 and Samuel, 1964) and equity prices as a gauge of long-run economic activity (Friedmann, 1988).

Non-stationarity is a condition for cointegration, so time series properties of the sample variables are examined by means of unit root tests. The ADF tests presented in Table 4.3. in Section 4.3.2. show that the variables employed in the adjusted hybrid model are first difference stationary for a significance level of 95%. The VAR for the adjusted hybrid

[21] According to U.S. Energy Information and Administration in 2012 Dubai Crude accounted for 83% of oil exports to Japan emanating from the Middle East Gulf. (https://www.eia.gov/todayinenergy/detail.php?id=13711). Dubai Crude is in addition a price benchmark for Persian Gulf exports to Asia (https://www.eia.gov/todayinenergy/detail.php?id=18571).

monetary model is grounded on the $X'_{(AdjRID)t}$ vector. Choice of lag length for the adjusted hybrid model is made by reference to model specification tests. A lag length of 4 is indicated for a well-specified VAR. The VAR specification for the adjusted hybrid model incorporates a constant, 3 seasonal dummy variables and impulse dummy variables that account for exogenous shocks pertaining to the USA saving and loan crisis in 1982 and 1989, Black Wednesday in 1992, the Asian financial crisis 1997-1998, the European Sovereign Debt Crisis 2010-2011 and collapse of the oil price at the start of 2015.

Misspecification tests for the adjusted hybrid model are reported in Appendix A4.2. (See Table A4.12.). The tests illustrate that the adjusted hybrid model of the USD/JPY exchange rate is free from serial correlation and autoregressive conditional heteroscedasticity at a significance level of 5%. The Portmanteau test shows residual autocorrelation in the productivity differential at the 5% significance level. Normality tests for the relative money supply, relative short-run interest rate, the real oil price and the system equation are significant at a level of 1% however, non-normality is rejected at a significance level of 1% for all other variables in the VAR. Since non-normality due to excess kurtosis does not affect cointegration (Gonzalo, 1994) we conclude from the tests of misspecification that the adjusted USD/JPY hybrid model is well formed.

Results of the Johansen (1995) test for cointegration for the adjusted hybrid monetary model of the USD/JPY are presented in Table 4.18. and show that the null hypothesis of no-cointegration is rejected, where there is confirmation of a single cointegrating vector at the 1% significance level. Evidence of a cointegrating vector in the adjusted USD/JPY hybrid model affirms that the ancillary variables for traded and non-traded goods productivity, asset prices and the real oil price observe the same stochastic trend as that of the nominal exchange rate and the monetary fundamentals included in the traditional RID monetary model. In sum, the adjusted hybrid monetary model of the USD/JPY exchange rate is a credible equilibrium model.

Table 4.18. Johansen (1995) Cointegration Test Results for the USD/JPY Adjusted Hybrid Monetary Model of the USD/JPY Exchange Rate The System comprises of [*e, m – m*, y – y, $i^s - i^{s*}$, $i^l - i^{l*}$, *s-s**, $prod^M$-$prod^{M*}$, *gs-gs** and *roil*]**						
(p –r)	*rank*	Eigenvalue	Trace Test	95% Critical Value	99% Critical Value	*p-value*
8	*r = 0*	0.46	246.75	192.89	204.95	[0.00] **
7	*r ≤ 1*	0.28	157.92	156.00	168.36	[0.06]
6	*r ≤ 2*	0.23	111.07	124.24	133.57	[0.27]
5	*r ≤ 3*	0.21	72.73	94.15	103.18	[0.63]
4	*r ≤ 4*	0.12	38.18	68.52	76.07	[0.97]
3	*r ≤ 5*	0.06	19.87	47.21	54.46	[1.00]
2	*r ≤ 6*	0.03	11.10	29.68	35.65	[0.95]
1	*r ≤ 7*	0.02	6.17	15.41	20.04	[0.68]
0	*r≤ 8*	0.02	2.55	3.76	6.65	[0.111]
Notes: *r* denotes the number of cointegrating vectors. ** indicates statistical significance at the 1% level. P-values in brackets [ ].						

Asymptotically distributed chi squared (Johansen, 1992b; Asteriou and Hall, 2015) tests of LE, WE and stationarity are presented in Table 4.19. and examine the extent to which the variables related to money, non-traded and traded goods productivity, real asset prices and the real oil price explain the nominal USD/JPY exchange rate in the adjusted framework. LE is examined by placing a zero restriction on the pertinent variable of β. Normalisation of the long-run relation on the variable of β is conditional upon rejection of the zero restriction. WE is investigated by means of a zero restriction on the pertinent variables of α. A variable is thought to be weakly exogenous if the zero restriction cannot be rejected inferring that the variable moves the long-run equilibrium instead of adapting to it. In a multivariate framework, the null hypothesis of stationarity is scrutinized by setting each variable in a single cointegrating vector to unity and the remaining variables to zero.

The stationarity tests (See Table 4.19., Panel C) in the adjusted hybrid model attest to the fact that all model variables are level nonstationary in the multivariate framework. LE tests for the adjusted hybrid monetary model of the USD/JPY exchange rate show that the nominal USD/JPY exchange rate can be excluded from the cointegrating framework. Differentials related to the discount interest rate and productivity in manufacturing as well as the real oil price can also be excluded from the adjusted long-run relation.

Table 4.19. LE, WE, and Stationarity (S) Tests for the Adjusted Hybrid Monetary Model of the USD/JPY Exchange Rate

	e	m-m*	y-y*	$i^s - i^{s*}$	$i^l - i^{l*}$	s-s*	$prod^M$-$prod^{M*}$	gs-gs*	r_oil
Panel A: LE tests									
$X^2(1)$	1.17	11.63	16.382	0.97	21.54	19.24	0.66	26.79	2.58
p-value	[0.28]	[0.00] **	[0.00] **	[0.32]	[0.00] **	[0.00] **	[0.42]	[0.00] **	[0.11]
Panel B: WE tests									
$X^2(1)$	2.46	2.99	0.60	1.10	17.89	6.40	0.38	4.33	6.79
p-value	[0.12]	[0.08]	[0.44]	[0.30]	[0.00] **	[0.01] *	[0.54]	[0.04] *	[0.01] **
Panel C: S tests									
$X^2(1)$	70.3	68.55	78.61	77.85	69.78	77.1	83.30	78.21	75.00
p-value	[0.00] **	[0.00] **	[0.00] **	[0.00] **	[0.00] **	[0.00] **	[0.00] **	[0.00] **	[0.00] **

Notes: ** and * indicate significance at the 1% and 5% levels, respectively. p-values are in brackets [].

Tests of WE for the variables in the adjusted hybrid monetary model show that WE is rejected at a significance level of 1% for the long-run interest rate differential and the real oil price. Subsequently, the term structure of interest rates is upheld since the relative long-run interest rate is endogenous to the cointegrating framework while the relative discount interest rate is weakly exogenous. This finding maintains that the direction of transmission is from the short-term interest rate to the long-term interest rate. Rejection of WE for the real oil price refutes the argument that the real oil price as exogenous to the US macroeconomic structure since shocks to the oil price originate from political unrest in the Middle East (Johansen and Juselius, 1992; Hamilton, 1985, Amano and van Norden; 1998; Kim and Roubini, 2000; Hunter and Ali, 2014). The inability to reject WE highlights the role of differentials linked to real asset prices and government consumption expenditure and indicates that shocks to the USD/JPY nominal exchange rate do indeed emanate real economic activity, long-run expectations and non-traded goods consumption (Rogoff, 1992; Lastrapes, 1992; Chen and Wu, 1997; Enders and Lee, 1997 and An and Kim, 2010).

Incorporation of variables related to traded and non-traded goods productivity, asset prices and the real oil price in the traditional RID monetary model has however not reversed the long run projection for the nominal USD/JPY exchange rate since the variable in question can be long-run excluded (see Table 4.19., Panel A). The nominal USD/JPY exchange rate is also shown to be weakly exogenous to the cointegrating relation estimated by the adjusted hybrid model (see Table 4.19., Panel B). The result of WE for the USD/JPY exchange rate derived in the RID model is not reversed in the adjusted model so cannot endorse this adaptation of the long-run evaluation.

An estimation of the long-run relation is not presented at this stage since normalisation is disqualified on variables that are either long-run excluded or weakly exogenous (Burke and Hunter, 2005, ch, 5). Joint tests of WE and LE are subsequently performed to derive a more coherent empirical model and adept identification of the long-run exchange rate equation.

4.4.4. Validation of the Adjusted Hybrid Model

WE and LE tests in section 4.4.3. accentuate the inadequacy of the adjusted hybrid monetary model as a long-run explanation of the USD/JPY exchange rate. In this section Hunter and Ali (2014) are followed to refine the adjusted model and establish a more compact yet powerful specification of the long-term exchange rate relation. The approach implemented is general-to-specific (Davidson et. al, 1978 and Hendry and Mizon, 1993) and founded on the outcome of LE and WE tests. A single cointegrating vector of rank r=1 is distinguished in the adjusted hybrid model of the USD/JPY exchange rate as well as the following framework referring to the α and β vectors:

	e	∇m	∇y	∇i^s	∇i^l	∇s	$\nabla prod^M$	∇gs	roil
$\beta' =$	[0	β_2	β_3	β_4	β_5	β_6	β_7	β_8	β_9],
$\beta' =$	[β_1	0	β_3	β_4	β_5	β_6	β_7	β_8	β_9],
$\beta' =$	[β_1	β_2	0	β_4	β_5	β_6	β_7	β_8	β_9],
$\beta' =$	[β_1	β_2	β_3	0	β_5	β_6	β_7	β_8	β_9],
$\beta' =$	[β_1	β_2	β_3	β_4	0	β_6	β_7	β_8	β_9],
$\beta' =$	[β_1	β_2	β_3	β_4	β_5	0	β_7	β_8	β_9],
$\beta' =$	[β_1	β_2	β_3	β_4	β_5	β_6	0	β_8	β_9],
$\beta' =$	[β_1	β_2	β_3	β_4	β_5	β_6	β_7	0	β_9] and
$\beta' =$	[β_1	β_2	β_3	β_4	β_5	β_6	β_7	β_8	0]

Further:

	e	∇m	∇y	∇i^s	∇i^l	∇s	$\nabla prod^M$	∇gs	roil
$\alpha' =$	[α_1	α_2	α_3	α_4	α_5	α_6	α_7	α_8	α_9]
$\alpha' =$	[0	α_2	α_3	α_4	α_5	α_6	α_7	α_8	α_9],
$\alpha' =$	[α_1	0	α_3	α_4	α_5	α_6	α_7	α_8	α_9],
$\alpha' =$	[α_1	α_2	0	α_4	α_5	α_6	α_7	α_8	α_9],
$\alpha' =$	[α_1	α_2	α_3	0	α_5	α_6	α_7	α_8	α_9],
$\alpha' =$	[α_1	α_2	α_3	α_4	0	α_6	α_7	α_8	α_9],
$\alpha' =$	[α_1	α_2	α_3	α_4	α_5	0	α_7	α_8	α_9] and
$\alpha' =$	[α_1	α_2	α_3	α_4	α_5	α_6	0	α_8	α_9]

Where ∇ represents the differential between domestic (USA) and foreign variables (Japan).

Based on this framework we impose sequentially the restriction on LE for the JPY/USD adjusted hybrid monetary model and sequentially impose zero restrictions on the loading factors, α, that are respectively weakly exogenous to the cointegrating relation (See Table 4.19., Panel B). The restrictions on weakly exogenous variables are empirically feasible in view of monetary doctrine and the magnitude of variable coefficients. In view of the individual LE and WE tests presented in Table 4.19, strict exogeneity (SE) is imposed on the variable relating to relative productivity with $\beta_7=0$ and $\alpha_7=0$ and on the variable relating to the relative short-run interest rate with $\beta_4=0$ and $\alpha_4=0$. Student's t-test statistics for the alpha coefficients of the subsequent restricted hybrid model with normalisation on the nominal exchange rate ($\beta_1=-1$) are presented in Table 4.20. and the significance, based on standard inference, of alpha coefficients is appraised to further refine the model specification.

Table 4.20. The Estimated Cointegrating Vector for the Adjusted Hybrid Model of the JPY/USD With Strict Exogeneity Imposed on Relative Productivity and the Relative Short-Run Interest Rate

	e	m-m*	y-y*	$i^l - i^{l*}$	s-s*	gs-gs*	roil
alpha	-0.01	-0.00	0.00	-0.14	0.02	0.04	-0.03
standard errors of alpha	0.01	0.00	0.00	0.033	0.01	0.02	0.01
t-stat	-1.68	-1.62	0.98	-4.19**	2.43*	2.05*	-2.5*

Notes: The critical values of the t-statistic are based on standard inference. ** and * indicate significance at the 1% and 5% levels, respectively.

In Table 4.20. it is shown that the alpha coefficients on relative money and relative income are insignificant. WE is as a consequence imposed on the insignificant variables of alpha with $\alpha_3=0$ on relative income and $\alpha_2=0$ on relative money. Joint tests of strict exogeneity (SE) and WE for the adjusted hybrid monetary model of the JPY/USD exchange rate are presented in Table 4.21. as follows:

Table 4.21. Exclusion of $prod^M$-$prod^{M*}$ and i^s-i^{s*} for the Adjusted Hybrid Model of the JPY/USD Exchange Rate						
Panel A: Joint Tests of LE and WE and Conditional on r = 1						
Tests under the null:						Statistics [p-value]
(1) β_7=0, α_7=0						X^2 (1) = 0.78 [0.68]
(2) β_7=0, α_7=0, B_4=0, α_4=0						X^2 (2) = 5.68 [0.22]
(3) β_7=0, α_7=0, B_4=0, α_4=0, α_3=0,						X^2 (3) = 6.90 [0.23]
(4) β_7=0, α_7=0, B_4=0, α_4=0, α_3=0, α_2=0						X^2 (4) = 10.1 [0.12]
Panel B: The Implied Long-Run Relation by Test (4):						
	(m-m*)	(y-y*)	(i^l-i^{l*})	(s-s*)	(gs-gs*)	(r_oil)
Coef.	-4.74	-34.25	1.08	-4.70	-2.62	1.05
Standard error	1.17	6.50	0.15	0.83	0.36	0.48
t-statistic	-4.07**	-5.27**	7.14**	-5.69**	-7.25**	2.19*
Notes: The critical values of the t-statistic are based on standard inference. ** and * indicate significance at the 1% and 5% levels, respectively. p-values are in square brackets[.].						

Results of the joint tests of LE and WE, presented in Table 4.21. show that the restrictions imposed on the variables in the JPY/USD adjusted hybrid model are firmly accepted for strict exogeneity of the differentials relating to the short-run interest rate (β_4=0 and α_4=0) and productivity (β_7=0 and α_7=0) as well as WE of the relative money supply (α_2=0) and relative income (α_3=0). The restricted long-term relationships, formerly discussed, normalised on the exchange rate show that all variables are significant, at the 1% level based on a standard inference, and all the variable coefficients apart from the relative money supply and real oil price possess hypothesised signs.

The implied long-run relation after imposing joint restrictions for LE and WE is illustrated in Table 4.21 (see Panel B). The exchange rate is denoted in terms of the domestic currency (USDs per JPY) so that a coefficient rise indicates a depreciation of the domestic currency. The coefficient on relative money is negative (β_1<0) and so prompts an appreciation of the domestic currency as compared to a theoretical one for one depreciation of the domestic currency. The coefficient on the income differential is negative (β_2<0) so confirms the hypothesis that a rise in domestic income relative to foreign income heightens the demand for the domestic currency and brings about an appreciation of the domestic currency. The coefficient on the long-term interest rate differential is positive (β_4>0) and supports the theory that a fall in the expected rate of domestic inflation stimulates capital inflows and creates increased demand for domestic currency relative to foreign and thereby prompts a domestic currency gain. Here a 1% rise in the long-run interest rate indicates a rise in anticipated inflation that inhibits capital inflows and the demand for domestic currency thereby prompting a depreciation of the domestic currency. However, when capital inflows increase when the domestic interest rate is raised thus causing an appreciation of the domestic currency. The coefficient on the relative share price is negative (β_5<0) so that the wealth effect governs money demand functions between the USA and Japan inducing an appreciation of the domestic currency (Friedman, 1988; McCornac, 1991; Caruso, 2001). The coefficient on relative government spending is negative (β_6<0) so supports the theory that an increase in government consumption expenditure creates a heightened demand for non-traded goods and as a result a price rise that prompts an appreciation of the domestic currency. The coefficient on oil is positive (β_7>0) so motivates a depreciation of the USD. This outcome opposes the argument that oil exports to Japan, in view of their enormity, should induce an appreciation of the USD since oil is priced in USDs (Johansen and Juselius, 1992; Amano and van Norden, 1998; Lizardo and Mollick, 2010 and Chinn and Moore, 2011).

These results empirically confirm stability of the adjusted hybrid monetary model as a model of the nominal JPY/USD exchange rate and the importance of variables linked to money, real share prices, non-traded goods consumption and the real oil price in determining the long-run relation.

4.5. Conclusions

In this paper, we review the USD vis-à-vis the JPY by application of the traded goods consumption smoothing model proposed by Rogoff (1992), the standard real interest differential model (Frankel, 1979a) and an adjusted version of the hybrid monetary model proposed by Hunter and Ali (2014). The adjusted version of (Hunter and Ali, 2014) is developed by incorporating variables related to traded and non-traded goods productivity (Rogoff, 1992), variables affecting stable money demand and PPP as well as real share prices that encompass facets of long-run economic expectations. A real oil price is in addition incorporated in the adjusted hybrid model to account for terms of trade. Quarterly data is employed from 1980:Q1 to 2016:Q4 a time characterized flexible exchange rate regimes, financial market integration and heightened capital mobility, the worst financial crisis since the great depression and its aftermath giving rise to a prolonged period of zero lower bound interest rates. The JPY/USD exchange rate models under examination are estimated by application of the Johansen (1995) methodology. A single long-run cointegrating relation is identified for the traded goods consumption smoothing, RID and adjusted hybrid monetary models of the JPY/USD exchange rates.

The traded goods consumption smoothing model is shown to be an incoherent exchange rate model as the variable of primary interest, the real JPY/USD^{CPI} exchange rate is found to be weakly exogenous. The standard RID monetary model yields a similar result as the explanatory variables cannot be normalised on the nominal exchange rate as it is found to be exogenous to the cointegrating system. The JPY/USD exchange rate in the traded goods consumption smoothing model and the standard RID model appears to behave like a random walk.

The adjusted hybrid monetary model is subsequently constructed by including variables employed in the traded goods consumption smoothing model as well as the standard RID monetary model and real asset prices. It is illustrated by means of joint tests of LE and WE that the adjusted hybrid monetary model produces a more articulate description of the identified long-run relation for the JPY/USD. The adjusted hybrid monetary model surpasses the traded goods consumption smoothing model as well as the standard RID model as a model of the JPY/USD exchange rate that is found to be endogenous to the system and cannot be long-run excluded. All variable coefficients are in addition found to be significant at the 1% level, based on standard inference. A shortfall of the adjusted hybrid model is that the real oil price is endogenous to the cointegrating relation and the variable coefficient is of a sign opposing the theory that a rise in the real price of oil prompts an appreciation of the USD. The real oil price therefore exhibits a powerful long-term effect that is relayed via long-term dynamics.

In sum model variables linked to monetary fundamentals have overturned their long-run projection for the JPY/USD exchange rate and it is observed that the JPY/USD is driven by the cointegrating framework in the adjusted hybrid monetary model and does not drive the framework portrayed in the traded goods consumption smoothing model or the standard RID model. It can therefore be concluded that variables related to monetary fundamentals, cross-country productivity differentials, long-run economic expectations and terms of trade are central to this model of the long-run JPY/USD exchange rate

Appendix A4.1.

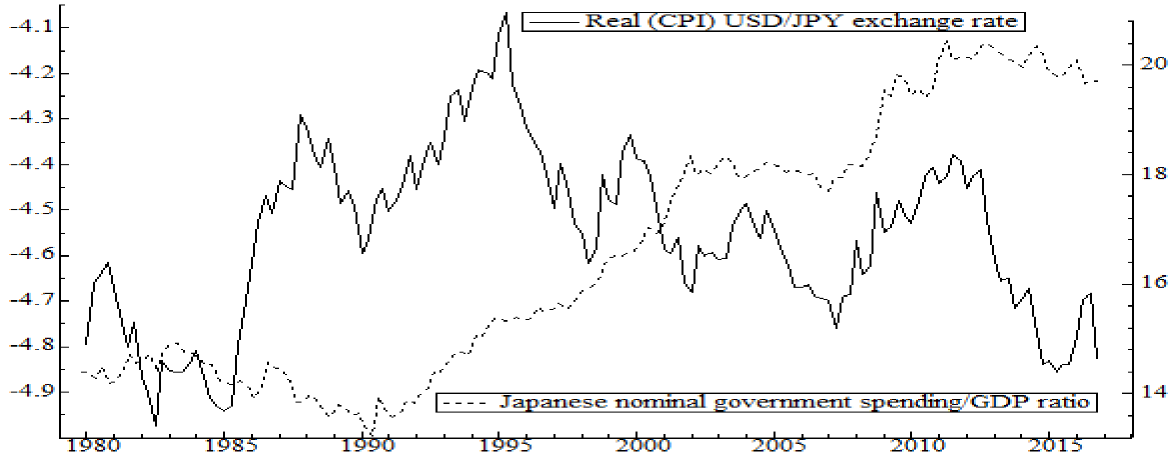

Figure A4.1. The Real (CPI) USD/JPY Exchange Rate (Left-Hand Scale) Versus The Nominal Japanese Government Spending/GDP Ratio (Right-Hand Scale)

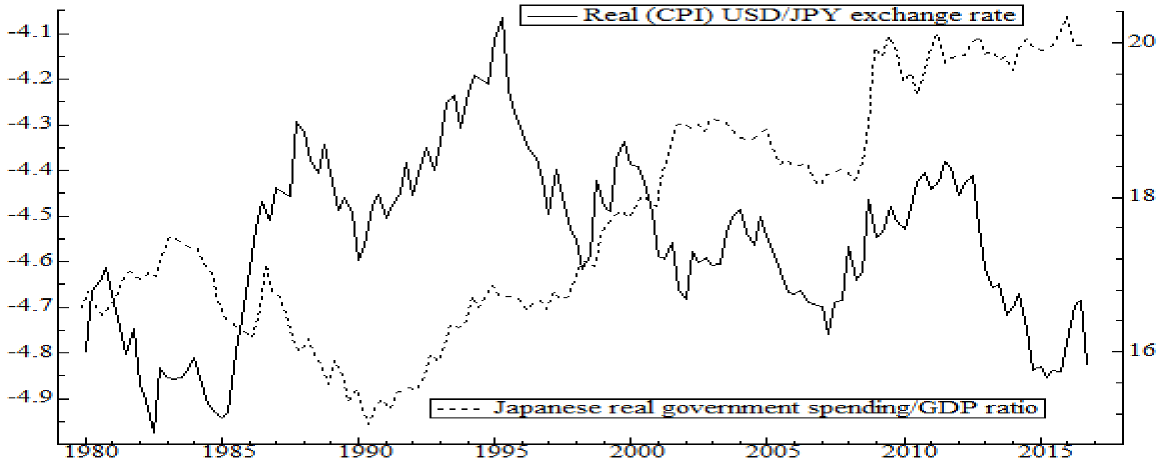

Figure A4.2. The Real (CPI) USD/JPY Exchange Rate (Left-Hand Scale) Versus the Real Japanese Government Spending/GDP Ratio (Right-Hand Scale)

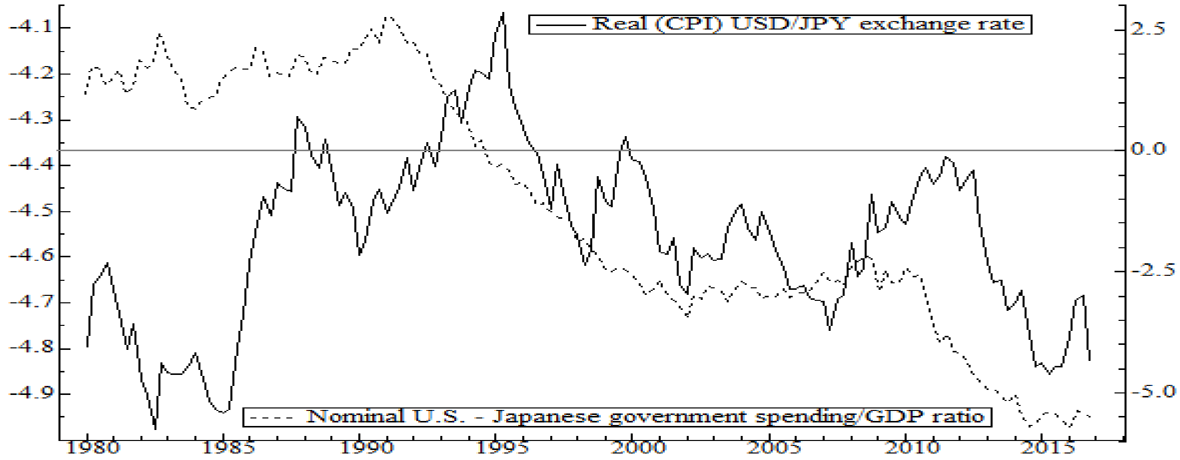

Figure A4.3. The Real (CPI) USD/JPY Exchange Rate (Left-Hand Scale) Versus the Nominal US - Japanese Government Spending/GDP Ratio (Right-Hand Scale)

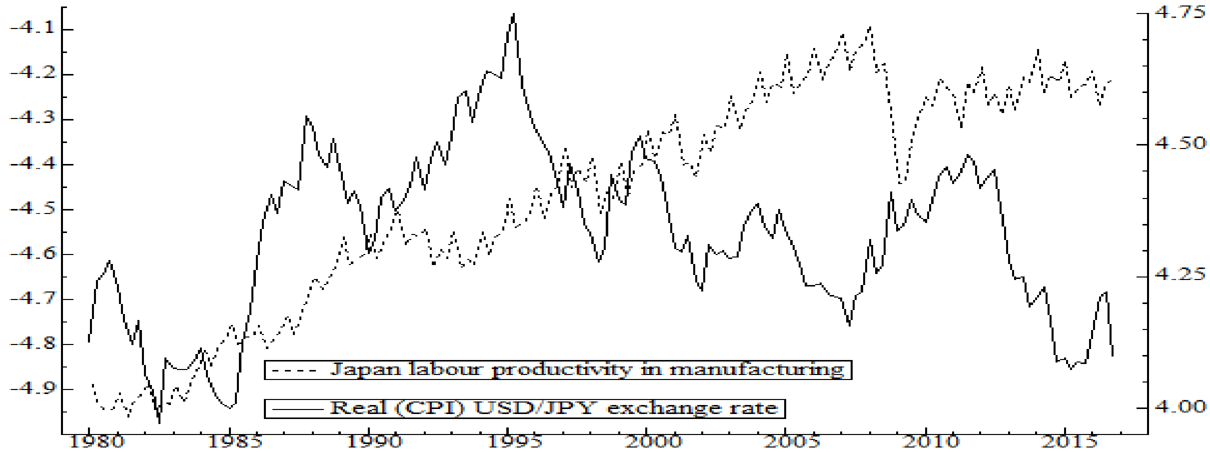

Figure A4.4. The Real (CPI) USD/JPY Exchange Rate (Left-Hand Scale) Versus Japanese Labour Productivity in Manufacturing (Right-Hand Scale)

Appendix A4.2.

Table A4.1. Misspecification Tests for the Traded and Non-Traded Goods Productivity Model of the Real USD/JPY^{CPI} Exchange Rate						
	Portmanteau(12)	LM(8)	ARCH(8)	Normality	Skewness	Excess Kurtosis
Panel A: Single-Equation Diagnostics Using Reduced-Form Residuals for Traded and Non-Traded Goods Productivity Model of the real USD/JPY^{CPI} Exchange Rate						
RER^{CPI}	8.81 [0.36]	1.29 [0.26]	0.78[0.62]	1.25 [0.54]	0.19	3.13
prod^{M}-prod^{M*}	45.35 [0.00]**	9.27 [0.00]**	2.03 [0.05]*	1.32 [0.52]	0.22	2.63
gs-gs*	11.38 [0.18]	1.3 [0.25]	0.81[0.6]	3.17 [0.21]	-0.22	3.62
roil	11.83 [0.16]	1.11 [0.36]	1.46[0.18]	6.87 [0.03]*	-0.30	3.99
Panel B: System Tests LM (8) test: 1.19 [0.11] Normality: 14.92 [0.06]						
Notes: LM (8) Is a Lagrange multiplier test of serial correlation up to order 8; p values are reported in square brackets [.].						

Table A4.2. Exclusion of roil for the Traded and Non-Traded Goods Productivity Model of the Real USD/JPY^{CPI} Exchange Rate		
Joint tests of WE and LE conditional on r = 1		
Tests under the null:		Statistics [p-value]
(1) B_4=0		X^2 (1) = 0.54 [0.46]
(2) B_4=0, α_4=0 oil is strictly exogenous		X^2 (2) = 1.48 [0.48]
(3) B_4=0, α_4=0, α_1=0		X^2 (3) = 3.63 [0.30]
The implied long-run relation by test (3):		
	prod^{M}-prod^{M*}	gs-gs*
Coef.	3.75	0.02
Standard error	0.58	0.01
t-stat	6.46**	1.5
Notes: ** and * indicate significance at the 1% and 5% levels, respectively based on standard inference. p-values are in square brackets[.].		

Table A4.3. Exclusion of gs-gs* for the Traded and Non-Traded Goods Productivity Model of the real USD/JPY^{CPI} Exchange Rate		
Joint tests of WE and LE conditional on r = 1		
Tests under the null:		Statistics [p-value]
(1) B_3=0		X^2 (1) = 2.26 [0.13]
(2) B_3=0, α_4=0		X^2 (2) = 2.66 [0.26]
(3) B_3=0, α_4=0, α_1=0		X^2 (3) = 3.91 [0.27]
The implied long-run relation by test (3):		
	prod^{M}-prod^{M*}	roil
Coef.	4.83	-0.16
Standard error	0.91	0.10
t-stat	5.36**	-1.58
Notes: ** and * indicate significance at the 1% and 5% levels, respectively based on standard inference. p-values are in square brackets[.].		

Table A4.4. Misspecification Tests for the Traded and Non-Traded Goods Productivity Model of the real USD/JPY^{WPI} Exchange Rate						
	Portmanteau(12)	LM(8)	ARCH(8)	Normality	Skewness	Excess Kurtosis
Panel A: Single-equation diagnostics using reduced-form residuals for traded and non-traded goods productivity model of the real USD/JPY^{WPI} exchange rate.						
RER^{WPI}	4.92 [0.77]	1.06 [0.39]	0.64 [0.75]	2.0 [0.37]	0.22	3.03
prod^{M}-prod^{M*}	58.37 [0.00]**	10.72 [0.00]**	1.44 [0.19]	1.4 [0.50]	0.15	2.6
gs-gs*	11.31 [0.18]	1.34 [0.23]	2.46 [0.02]*	1.85 [0.4]	-0.07	3.4
roil	12.52 [0.13]	1.48 [0.17]	1.76 [0.09]	5.62[0.06]	-0.28	3.82
Panel B: System Tests LM (8) test: 1.28 [0.04]* Normality: 10.40 [0.24]						
Notes: LM (8) Is a Lagrange multiplier test of serial correlation up to order 8; p values are reported in square brackets [.].						

Table A4.5. The Estimated Cointegrating Vector for the Traded and Non-Traded Goods Productivity model normalised on the real USD/JPY^{WPI} exchange rate			
	prod^{M}-prod^{M*}	gs-gs*	roil
Coefficient	3.74	-0.01	-0.02
Standard error	0.83	0.02	0.10
t-stat	4.48**	-0.6	-0.16
Notes: ** and * indicate significance at the 1% and 5% levels, respectively based on standard inference. p-values are in square brackets[.].			

Table A4.6. Exclusion of roil for the traded and non-traded goods productivity model of the real USD/JPY^{WPI} Exchange Rate		
Joint tests of WE and LE conditional on r = 1		
Tests under the null:		Statistics [p-value]
(1) B_4=0		X^2 (1) = 0.01 [0.93]
(2) B_4=0, α_4=0		X^2 (2) = 0.03 [0.99]
(3) B_4=0, α_4=0, α_1=0		X^2 (3) = 1.82 [0.61]
(4) B_4=0, α_4=0, α_1=0, α_3=0		X^2 (4) = 4.34 [0.36]
The implied long-run relation by test (4):		
	prod^{M}-prod^{M*}	gs-gs*
Coef.	4.30	0.01
Standard error	0.76	0.02
t-stat	5.72**	0.25
Notes: ** and * indicate significance at the 1% and 5% levels, respectively based on standard inference. p-values are in square brackets[.].		

Table A4.7. Exclusion of gs-gs* for the Traded and Non-Traded Goods Productivity Model of the real USD/JPY^{WPI} Exchange Rate	
Joint tests of WE and LE conditional on r = 1	
Tests under the null:	Statistics [p-value]
(1) B_3=0	X^2(1) = 0.16 [0.69]
(2) B_3=0, α_3=0	X^2(2) = 2.57 [0.28]
(3) B_3=0, α_3=0, α_4=0	X^2(3) = 2.61 [0.46]
(4) B_3=0, α_3=0, α_4=0, α_1=0	X^2(4) = 2.67 [0.61]

The implied long-run relation by test (4):

	$prod^M$-$prod^{M*}$	roil
Coef.	6.43	-0.24
Standard error	1.33	0.15
t-stat	4.82**	-1.57

Notes: ** and * indicate significance at the 1% and 5% levels, respectively based on standard inference. p-values are in square brackets[.].

Table A4.8. Misspecification Tests of the RID Model of the USD/JPY Exchange Rate

	Portmanteau(12)	LM(8)	ARCH(8)	Normality	Skewness	Excess Kurtosis
Panel A: Single-equation diagnostics for the USD/JPY RID Model						
e	13.27 [0.10]	0.98 [0.46]	0.46 [0.88]	2.58 [0.28]	0.34	3.12
m-m*	9.42 [0.31]	1.14 [0.34]	0.89 [0.53]	5.52 [0.06]	-0.12	3.52
y-y*	9.43 [0.31]	0.53 [0.83]	0.94 [0.49]	3.19 [0.20]	-0.16	3.39
$i^s - i^{s*}$	9.65 [0.29]	0.92 [0.50]	1.82 [0.08]	10.42 [0.01]**	-0.68	3.63
$i^l - i^{l*}$	6.79 [0.56]	0.53 [0.83]	1.16 [0.33]	2.82 [0.24]	0.34	3.44

Panel B: System Tests
LM (8) test: 1.30 [0.02]*
Normality: 22.96 [0.01]*

Notes: LM (8) Is a Lagrange multiplier test of serial correlation up to order 8; p values are reported in square brackets [.]. * and ** Indicate statistical significance at the 1% and 5% level, respectively.

Table A4.9. Exclusion of y-y* for the RID Model of the USD/JPY Exchange Rate	
Joint tests of WE and LE conditional on r = 1 in the HM model	
Tests under the null:	Statistics [p-value]
(1) B_3=0	X^2(1) = 3.00 [0.58]
(2) B_3=0, α_3=0	X^2(2) = 0.44 [0.80]
(3) B_3=0, α_3=0, α_1=0	X^2(3) = 3.48 [0.32]

The implied long-run relation by test (3):

	(m-m*)	(i^s-i^s*)	(i^l-i^l*)	
Coef.	0.30	0.07	0.28	
Standard error	0.16	0.04	0.06	
t-stat	1.86*	1.74*	4.42**	

Notes: ** and * indicate significance at the 1% and 5% levels, respectively based on standard inference. p-values are in square brackets[.].

Table A4.10. Exclusion of m-m* for the RID Model of the USD/JPY Exchange Rate				
Joint tests of WE and LE conditional on r = 1 in the HM model				
Tests under the null:				Statistics [p-value]
(1) B_2=0				X^2(1) = 2.77 [0.10]
(2) B_2=0, α_3=0				X^2(2) = 3.06 [0.22]
(3) B_2=0, α_3=0, α_1=0				X^2(3) = 5.08 [0.17]
The implied long-run relation by test (3):				
	(y-y*)	(i^s-i^s*)	(i^l-i^l*)	
Coef.	-0.35	0.03	0.34	
Standard error	0.39	0.04	0.06	
t-stat	-0.92	0.70	5.27**	
Notes: ** and * indicate significance at the 1% and 5% levels, respectively based on standard inference. p-values are in square brackets[.].				

Table A4.11. Exclusion of i^s-i^{s*} for the RID Model of the USD/JPY Exchange Rate				
Joint tests of WE and LE conditional on r = 1 in the HM model				
Tests under the null:				Statistics [p-value]
(1) B_4=0				X^2(1) = 3.24 [0.07]
(2) B_4=0, α_3=0				X^2(2) = 3.39 [0.18]
(3) B_4=0, α_3=0, α_1=0				X^2(3) = 4.90 [0.18]
The implied long-run relation by test (3):				
	(m-m*)	(y-y*)	(i^l-i^l*)	
Coef.	0.15	0.05	0.38	
Standard error	0.23	0.58	0.04	
t-stat	0.64	0.01	8.81**	
Notes: ** and * indicate significance at the 1% and 5% levels, respectively based on standard inference. p-values are in square brackets[.].				

Table A4.12. Misspecification Tests of the Hybrid Model for the USD/JPY Exchange Rate						
	Portmanteau(12)	LM(8)	ARCH(8)	Normality	Skewness	Excess Kurtosis
Panel A: Single equation tests						
e	11.96 [0.15]	1.02 [0.42]	0.35 [0.95]	3.56 [0.17]	0.33	3.55
$m - m^*$	14.91 [0.06]	2.06 [0.05]*	0.46 [0.89]	69.88 [0.00]**	-1.58	12.73
$y - y^*$	13.49 [0.10]	1.7 [0.11]	0.59 [0.79]	1.92 [0.38]	0.05	3.29
$i^s - i^{s*}$	18.11 [0.02]*	1.8 [0.09]	2.11 [0.04]*	12.84 [0.00]**	-0.60	3.67
$i^l - i^{l*}$	6.26 [0.62]	1.42 [0.2]	0.64 [0.74]	6.55 [0.04]*	-0.18	3.66
s-s^*	11.71 [0.17]	2.22 [0.03]*	1.9 [0.07]	2.83 [0.24]	0.32	2.98
$prod^M$-$prod^{M*}$	21.12 [0.01]**	2.24 [0.03]*	0.68 [0.71]	6.71 [0.04]*	-0.41	4.06
gs-gs^*	14.6 [0.07]	2.6 [0.01]*	0.4 [0.92]	1.81 [0.40]	-0.04	3.42
roil	15.30 [0.05]	2.01 [0.06]	0.9 [0.53]	15.13 [0.00]**	-0.01	4.61
Panel B. System Tests						
LM (8): 1.34 [0.03]* Normality: 108.11 [0.00]**						
Notes: LM (8) Is a Lagrange multiplier test of serial correlation up to order 8; p-values are reported in square brackets [.].** and * indicate significance at the 1% and 5% levels, respectively. p-values are in square brackets[.].						

CHAPTER FIVE

CONCLUSIONS

In this thesis a contribution is made to the extant literature pertaining to exchange rates and international finance. In the **second chapter** we revisit the standard RID monetary model (Frankel, 1979a) of the exchange rate that has universally failed to identify a coherent long-run relationship between exchange rates and monetary fundamentals. We progress from a study of the traditional monetary model to an examination of the hybrid monetary model developed by Hunter and Ali (2014) for the USD/JPY exchange rate. The hybrid model is employed to investigate the relationship between the GBP and CAD vis-à-vis and the USD and variables related to the monetary base and long-run economic expectations. The period under investigation is characterized by the aftermath of the greatest financial crisis since the Great Depression and for comparison with Japan is characterized by zero-lower bound interest rates, a phenomenon first experienced by Japan for a prolonged period of time. A unique combination of data related to monetary fundamentals, share prices, productivity differentials and terms of trade is employed in a hybrid monetary model. The Johansen (1995) methodology is applied to both the RID and hybrid model estimation and a single cointegrating relationship is identified for both models. However, tests of LE and WE indicate that the hybrid model produces a more coherent explanation of the long-run exchange rate in both cases. More specifically it is shown that the nominal CAD/USD exchange rate has, in the framework of the hybrid model, a reversed long-run projection so is driven by the system as compared to driving it. In sum the hybrid model outclasses the RID model since the long-run underlying fundamentals that drive the GBP/USD and CAD/USD exchange rates are accounted for by the incorporation of real stock prices that react to one another in global financial markets (Friedman, 1988).

In Chapter 2 we contribute to the extant literature by improving the long-run projection of the standard RID monetary model for the GBP and CAD vis-à-vis the USD in the era of flexible exchange rate regimes and zero lower bond interest rates by incorporating variables related to long-run expectations and real economic activity. It is evident from the hybrid monetary model that differentials related to share prices, government spending and productivity as well as a real global oil price have an impact on stable money demand and exchange rate persistence. A recommendation for further study of the models examined in chapter 2 is to test performance of the random walk benchmark in out-of-sample forecasting at different terms.

In the **third chapter** we test efficiency of the market for foreign currency by modelling the pricing relationship between spot and futures rates for five currency pairs. A panel of spot and futures rates for the GBP, CAD, AUD, BRL and ZAR vis-à-vis the USD is constructed at several intervals prior to expiry to test the hypothesis that the spot market is led by the market for futures. By evaluating the lead-lag pricing relationship and deviations from theoretical futures values we are able to assess the market's ability to internalize information from market agents who behave in accordance with information and outlook about the future. The stochastic properties of all five spot and futures rates are examined as stationarity is a condition for cointegration and by application of the Johansen (1995) test for cointegration a linear relationship is established for all currency pairs at intervals prior to maturity and at maturity. Unit root properties of the basis are furthermore investigated at the same intervals prior to expiry and at expiry and confirm that the basis for all five currency pairs is first difference stationary. Results of this study show that the spot and futures market for the GBP, CAD, AUD, BRL and ZAR vis-à-vis the USD is efficient so that the futures market enables risk free arbitrage opportunities and the spot market accomplishes its' role as an appropriator of means so that prices return to equilibrium levels.

In chapter 3 we contribute to existing studies by finding that the market for foreign currency is efficient in the way in which information is internalized and that the existence of arbitrage opportunities re-establish equilibrium levels. The majority of work in this area focuses on the 'majors' or currencies with an extended market share. We have however, examined two dominant emerging market currencies and observe scope for developing further studies pertaining to emerging market currencies that at present have a growing market share. We further understand of the grwing market for foreign currency that since the breakdown of the floating exchange rate regime adds another channel for systematic risk as well as reason for increased regulation in the derivatives market.

In the **fourth chapter** we re-examine the hybrid monetary model of the USD/JPY nominal exchange rate formulated by Hunter and Ali (2014) but adjust the model by incorporating the instrumental variables employed by Rogoff (1992) in a consumption smoothing fixed factor model of the USD/JPY real exchange rate. The USD/JPY exchange rate is revisited in view of the global economic crisis of 2007-2009 and sluggish recovery of the Japanese economy as compared with the USA. Gross domestic product in the USA returned to pre-crisis levels approximately two years following the mortgage backed security crisis as opposed to GDP in Japan that as of 2015:Q4 is recorded to be below pre-crisis levels. In view of the Japanese economy's sluggish recovery (Lo and Rogoff, 2015) we consider the HBS effect and follow Rogoff (1992) by evaluating the JPY/USD real exchange rate in terms of productivity differentials in the traded and non-traded productivity sectors. The relationship between the real USD/JPY exchange rate and differentials related to productivity in manufacturing, government consumption expenditure and the real oil price is examined by means of the Johansen (1995) test for cointegration. Tests of cointegration confirm a relationship between the real USD/JPY exchange rate and the variables related to traded and non-traded goods productivity. The USD/JPY exchange rate is further examined in the RID monetary model which fails to establish a coherent long-run relationship between the USD/JPY exchange rate and model variables as the nominal exchange is found to be weakly exogenous so behaves like a random walk. The study is furthered by examining the hybrid monetary model adjusted for traded and non-traded goods productivity differentials. The hybrid model produces a single cointegrating relationship and joint tests of LE and WE confirm that the cointegrating framework drives the nominal USD/JPY exchange rate. The long-run projection of the nominal USD/JPY exchange rate is reversed by considering productivity differentials in the sectors for traded and non-traded goods productivity.

We contribute to the existing literature in the third study by improving the outlook for the standard RID model of the USD/JPY nominal exchange rate. We focus on productivity differentials in view of the 2007-2009 financial crisis and the Japanese Economy's sluggish return to pre-crisis GDP levels. We find that incorporation of variables related to productivity in different sectors improves the long-run projection for the USD/JPY nominal exchange rate. A recommendation for future research is to study exchange rate determination and productivity differentials across diverse sectors (De Gregorio and Wolf, 1994) in the post 2007-2009 financial crisis period.

In sum we show that the traditional RID monetary model of the exchange for the USD vis-à-vis the GBP, CAD and JPY can be rejected. We find that a single econometric specification can be adapted to explain the long-run GBP/USD exchange rate while an extended model is effective in providing an explanation of the long run CAD/USD exchange rate. We also demonstrate that the hybrid monetary model of the USD/JPY exchange rate due to Hunter and Ali (2014) appears a specific case as the USD/JPY exchange rate in our model is not readily distinguished from a random walk in the context of a monetary model that considers traded and non-traded goods productivity differentials.

Considering findings in the chapter 2 we resolve that PPP is no longer a crucial driver of the exchange rate as financial assets play a significant role in driving the system and micro as well as macro-economic phenomena cannot be categorized by short-term and long-term interest rates in a globalised market. Study of the extended monetary models presented in chapters 2 and 4 can be extended by constructing a dynamic error correction model and by assessing out-of-sample forecasting performance.

In chapter 3 we contribute to existing studies by extending the analysis of Antoniou and Garrett (1993) with respect to futures contracts on the FTSE to the functioning of the spot and futures market for foreign currency. We effectively pool a matched panel of spot and futures contracts for the USD vis-à-vis the GBP, CAD, AUD, BRL and ZAR and show by means of cointegration analysis that the spot and futures market for the five bilateral exchange rates in our study functions effectively across developed and developing countries. We confirm that the spot and futures markets for foreign exchange function, for both traditional contract lengths in developed economies and shorter cases for developing markets, as one. This study is significant in view of the growing share of foreign exchange market transactions and the need for greater market transparency that would allow regulators to lay down the most appropriate directives. We show that there is scope for studying emerging market currencies considering a larger market share and accessibility of more coherent data. A recommendation for further study is to extend the model to other financial markets and to examine the notion that trading in the futures market heightens volatility in the spot market.

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