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The portfolio balance effect: an empirical analysis for Japan and Eurozone

Abstract

In this study the validity of the portfolio balance model in the short and in the long run for the Yen-U.S. dollar and Euro- U.S. dollar exchange rate is examined which is based on the Branson and Henderson specification. A distinguishing feature of the portfolio balance model among exchange rate models is the assumption of imperfect substitutability between domestic and foreign assets. The econometric method used here is the dynamic OLS approach which corrects for regressor endogeneity and is a robust method implemented in small samples. Furthermore, the stationarity of the variables is examined by unit root and stationary tests.

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1. Introduction

Most theoretical work on exchange rates in the 1970s assumed that exchange rates were primarily determined by equilibrium conditions in the markets for the stocks of domestic and foreign assets (money and government bonds), known as the "asset market approach." These models were not originally formulated in terms of utility-maximizing behavior, but subsequent work has explored the consistency of their hypothesized behavioral relations with maximizing theory. The simplest versions of the asset market models, called monetary models, have proved useful in explaining currency movements during high inflation periods. Experience with floating rates between the industrialized countries in the moderate inflation environment of the post-Bretton Woods era suggests that monetary shocks alone are not enough to explain exchange rates.

Work on asset market approaches to exchange rates falls into three categories: standard monetary models, monetary models with sticky prices and portfolio balance models. Standard monetary models determine the exchange rate with monetary equilibrium conditions and purchasing power parity. The monetary model predicts that money supply movements lead immediately to equal movements in the exchange rate. With interest rate parity and rational expectations this continues to be true, and the exchange rate is a random walk if either money or income are random walks or if money demand is interest inelastic. Monetary models with sticky prices maintain monetary neutrality across steady states but allow money to influence real variables in the short run. In the monetary models, it does not matter how the change in the money stock is created, the effect on the exchange rate is the same regardless of the source of money creation. In the sticky price models of Dornbusch and Frankel the exchange rate overshoots: monetary shocks cause more than proportionate changes in the exchange rate. In Driskill's stock/flow model overshooting may disappear if foreign and domestic assets are imperfect substitutes. Portfolio balance models emphasize wealth effects on asset demands and the role of the exchange rate in the valuation of foreign assets. Some versions also introduce the current account balance in its role of allocating wealth between countries. Unlike the monetary models, in which money affects the exchange rate through prices, at the portfolio balance model the exchange rate has a direct effect on asset demands through the valuation of foreign-denominated assets. If the net foreign asset position is negative some of the portfolio balance models predict that monetary expansion causes appreciation of the domestic currency.

Relative to the monetary models of exchange rate determination, the key modification of the portfolio model is the assumption that domestic and foreign securities are not perfect substitutes. The result is that a risk premium intrudes on the uncovered interest parity condition and supplies of bonds and other non-monetary assets intrude on the equation of exchange rate determination. The exchange rate is determined by the supply and demand for all foreign and domestic assets, not just by the supply and demand for money as in the monetary approach.

During the 1970s and 1980s there has been a thorough reworking of macroeconomic theory for open economies using a portfolio balance approach. According to this approach, equilibrium in financial markets occurs when the available stocks of national moneys and other financial assets are equal to the stock demands for these assets based on current wealth, and wealth accumulation continues only until current wealth is equal to desired wealth. Although portfolio balance models were originally developed to study movements of financial capital, variations in interest rates, and changes in stocks of international reserves under fixed exchange rates, they were quickly adapted to study movements of financial capital, variations in interest rates, and changes in the exchange rate under flexible exchange rates. The builders of portfolio balance models have not denied the desirability of deriving asset demands from explicit utility maximizing behavior. Indeed, they have attempted to establish the plausibility of their asset demands by appealing to microeconomic theory (in the case of non-monetary assets to the theory of portfolio selection and in the case of monetary assets to the theory of money demand).

The portfolio balance theory says that investors diversify their holdings among domestic and foreign assets (including bonds, if we do not rule them out a priori on the grounds of Ricardian equivalence) as functions of expected rates of return. The arbitrage opportunity which is derived from the expected return helps to determine exchange rates. Portfolio balance models imply that exchange rates, jointly with interest rates, result from the equilibrium of supply and demand for domestic and foreign assets, where these assets are allowed to be imperfect substitutes for each other. Dynamic adjustment of the exchange rate over time results from the fact that current account surpluses (deficits) correspond to accumulation (decumulation) of foreign assets, and that the current account itself depends on both the exchange rate and the stock of foreign assets.

2. Literature review

Since the mid 1970s there has developed a voluminous literature on the determination of exchange rates that has been influenced by the rational expectations literature that took off at the same time. There was disagreement in the literature over what are the relevant economic fundamentals and this has led to a variety of competing exchange rate models. The most well known models are the flexible price monetary models (Bilson (1978), Frenkel (1976), Mussa (1976)), the sticky price monetary models (Dornbusch (1976) and Frankel (1979)) and the portfolio balance models (Kouri (1976), Branson (1976, 1984)). In the monetary models, domestic and foreign bonds are assumed to be perfect substitutes with the uncovered interest rate parity condition holding continuously. The relevant economic fundamentals for exchange rate determination are considered to be the relative supply and demands for national money stocks. The difference between the two monetary models is that the flexible monetary model considers that goods prices are flexible in the short run, while the sticky price model assumes that goods prices only fully adjust over the longer run. In the flexible price monetary model, the exchange rate follows a purchasing power parity path, while in the sticky price monetary models the exchange rate can deviate in the short run from purchasing power parity. In the portfolio balance models, domestic and foreign bonds are imperfect substitutes and the existence of a risk premium means that the uncovered interest parity condition does not generally hold. Although these models have represented a significant theoretical advance, empirical testing of these theories has been unrewarding in Branson, Halttunen and Masson (1977), Hacche and Townend (1981), Bisignano and Hoover (1982) Dooley and Isard (1982), Meese and Rogoff (l983a and 1983b), Frankel (1984) and Backus(1984). The estimated coefficients are often insignificant and there is a persistent problem of residual autocorrelation. The portfolio balance model of exchange rates has not attracted a large empirical literature relative to the monetary class of models perhaps because the choice of non-monetary assets to be considered is difficult and data are not always available on a bilateral basis. Intertemporal portfolio balance models begin with the implicit assumption that asset markets are always in equilibrium [Boyer (1978), Masson (1981), Henderson and Rogoff (1981)]. Their dynamics are based upon a sequence of shortrun comparative statics results. However, these comparative statics results are valid only if exchange rates and interest rates move rapidly to whatever levels will bring about short-run equilibrium. The market forces that induce exchange rates to move in this fashion must be explained. Thus, any analysis of intertemporal stability should begin with the logically prior question of short-run Walrasian stability as in Cadsby (1987). Two types of test have been conducted in the relevant empirical literature. The first type is based on the reduced-form solution of the short-run portfolio balance model in order to measure its explanatory power, under the assumption that expectations are static. The second type of test concentrates on solving the portfolio balance model for the risk premium and testing for perfect substitutability of bonds denominated in different currencies: the inverted demand approach.

Mussa (1979) argued that observed changes in exchange rates have been predominantly unexpected as he fails to explain the major portion of observed changes in exchange rates. Unexpected changes in observed exchange rates can be attributed to revisions in expectations about future exchange rates in response to new information about the prospective future time paths of variables on which exchange rate expectations are based. To the extent that observed changes in exchange rates are predominantly unexpected, researchers cannot hope to predict them with much accuracy ex ante unless they have relatively advanced information about the variables on which exchange rate expectations are based. But it remains plausible that the unexpected component of exchange rate changes can be explained accurately ex post if researchers can accurately measure the magnitude and timing of the revisions in expectations about future asset stocks and wealth variables that underlie the magnitude and timing of revisions in expectations about future exchange rates. In the literature on exchange rate dynamics it is popular to assume, following Dornbusch

(1976) and Kouri (1976) that the system evolves toward a long-run steady state in which the real exchange rate reaches an equilibrium level that is consistent with balance in the trade or current accounts. In the short run, current accounts can exhibit substantial unexpected shifts, generating a variety of explanations and puzzles, and in many cases leading to revisions in expectations about the long-run path of the current account that would be consistent with any given path of the real exchange rate. Accordingly, it can be argued that rational economic agents should respond to the latter types of current amount surprises by revising their expectations of the long-run real exchange rate (the real exchange rate is consistent with long-run current account balance).Thus, unexpected shifts in current accounts that are perceived to be permanent, other things equal, can lead to revisions in expectations about future exchange rates and associated unexpected shifts in observed exchange rates.

The portfolio balance framework has attracted attention for emphasizing that international transfers of wealth through current account imbalances can influence exchange rates if assets denominated in different currencies are imperfect substitutes (Kouri (1976) and Branson (1977)). The literature that has explored this channel of influence, however, has developed almost uniformly around the assumption that current account imbalances are financed by transferring assets denominated in one of the two available currencies, while assets denominated in the other currency are not traded internationally. In models that adopt this assumption changes in the currency composition of asset portfolios can only occur through current account imbalances. This assumption is obviously inconsistent with the present functioning of international credit markets in which governments and private borrowers routinely issue foreign currency denominated debt. Branson (1983) uses a rational expectation version of the model of Kouri (1978) and shows how unanticipated movements in money, the current account, and relative price levels will cause first a jump in the exchange rate, and then a movement along a saddle path to the new long run equilibrium. Here the role of "news" in moving the exchange rate, as emphasized by Dornbusch (1980) and Frenkel (1981), is clear.

A major strand in the portfolio balance literature involves estimating log-linear versions of reduced form equations with the exchange rate as the dependent variable. For example, Artus (1976) and Branson, Halttunen and Masson (1977) study the mark/dollar rate while Driskill (1981) analyzes the Swiss franc/dollar rate. These studies typically report results supportive of the portfolio balance model. But in all cases, this support is based primarily on the explanatory power in exchange rate equations of either net private capital flows or the current account. The conditions for stability of a portfolio balance model of exchange rate determination with an endogenous current account are examined by Masson (1981) for various expectational assumptions. He shows that unless strongly stabilizing expectational assumptions are made, if the economy is a net debtor in foreign currency assets there is the possibility that its exchange rate will exhibit instability. It is true that, to the extent current account surpluses and deficits reflect differences in national savings rates, the portfolio balance model predicts that a current account surplus will be accompanied by an appreciating exchange rate. Holding the supply of outside assets constant, the exchange rate appreciates as the relative real wealth of a country's citizens rises, as they prefer assets denominated in their own currency.

Extensive surveys of models of international return differentials are found in Branson and Henderson (1985) and Adler and Dumas (1983). The optimal portfolio among foreign and domestic currency assets (and thus return differentials given supplies of home and foreign assets) depends on the covariance matrix between all asset returns, national outputs, national consumption levels, and national inflation rates. The empirical success of such portfolio-based models of international asset returns has not been encouraging (Frankel, 1982 and Boothe and Longworth, 1985).

While other studies limit the portfolio choice to domestic assets relative to a composite foreign asset, Lewis' approach (1988) is interesting because it considers a decomposition of the foreign asset by currency. Further, Lewis exploits the cross-equation correlation that arises from this decomposition in order to obtain more efficient estimates of the parameters, and provides some empirical support for the portfolio balance model.

There exists a variety of tests, both direct and indirect, of portfolio balance model validity. Direct reduced form tests of the portfolio balance model have been conducted by Branson, Haltunnen and Masson (1977), Dooley and Isard (1982), and by Hooper and Morton (1982). Hooper and Morton (1982) try to implement the portfolio balance theory within the Federal Reserve Board's Multi-Country Model, a large-scale econometric model consisting of quarterly models of Canada, Germany, Japan, the United Kingdom and the United States as well as abbreviated OPEC and rest-of-world sectors. Hooper and Morton explore numerous approaches, involving alternative econometric techniques, simplifying assumptions, data sets, and hypotheses concerning the expectations. The authors conclude that they are unable to find empirical evidence to confirm that the portfolio balance approach has quantitative significance. In addition, indirect tests have attempted to determine whether risk constitutes an important variable in foreign exchange markets so researchers typically estimate an equation where the risk premium is a function of domestic and foreign bond holdings. The imperfect substitutability of domestic and foreign assets which is assumed in the portfolio balance model is equivalent to assuming that there is a risk premium separating expected depreciation and this risk premium will be a function of relative domestic and foreign debt outstanding. The risk premium is measured by deviations from uncovered interest parity, either assuming rational expectations or employing survey data. In general, the empirical literature on testing the portfolio balance model suggests that sterilised intervention is effective at most in the very short term (Frankel, 1982a; Rogoff, 1984; Lewis, 1988; Edison, 1993), while the joint hypothesis of rational expectations and perfect substitutability of domestic and foreign assets is regularly rejected. Much of this literature suggests that the exchange rate effects of intervention through the portfolio balance channel are very small in size (Frankel, 1982a; Obstfeld, 1983; Rogoff, 1984; Danker, Haas, Henderson, Symansky and Tryon, 1987; Lewis, 1988). For example, if assets are perfect substitutes, capital perfectly mobile and expectations rational, the risk premium should be serially uncorrelated, orthogonal to the information set on which expectations are based and regressing it on variables determining risk should result in insignificant coefficients. Danker, Haas, Henderson, Symansky and Tryon (1987) attempt to detect a portfolio balance effect in separate bilateral equations for the mark/dollar, yen/dollar, and Canadian dollar/U.S. Dollar uncovered interest rate differentials. Danker et al. derive their risk premium equations by inverting bond demand equations which are disaggregated between the bank and non-bank private sectors. Despite these refinements, Danker et al.'s estimated risk premium equations (estimated under both static and rational expectations) provide little evidence in support of their model. The portfolio balance variables are jointly and individuallly insignificant for the Canadian dollar/US. dollar and yen/dollar equations. Although the portfolio balance variables are jointly significant in the mark/dollar equations, many of the individual coefficients are of the wrong sign. Danker et al. also report the results of an extensive specification search, in which they succeed in obtaining results for Germany and Japan (but not for Canada) which broadly conform to the theoretical predictions of the portfolio balance model. Frankel (1982) using monthly data for a selection of currencies regresses risk premium on a selection of the determinants of risk and reports a statistically insignificant relationship. Frankel exploits the fact that the coefficients of inverted portfolio balance equations are found to be related to the variance-covariance matrix. He then estimates a portfolio balance model with mean-variance optimization and tests the null hypothesis that the parameter representing the coefficient of risk aversion is equal to zero against the alternative hypothesis that rates of return are related to asset supplies in the portfolio balance model. His empirical results suggest that no statistical link appears to exist between asset supplies and the risk premium, implying that sterilised intervention cannot influence the risk premium or the exchange rate and is therefore ineffective. Frankel (1982a) assumes that market participants have rational expectations and constructs the 'ex-post' uncovered interest differential by using the actual rate of exchange rate appreciation in place of its unobservable expected value. For explanatory variables, he employs portfolio balance variables such as the relative supplies of central government bonds (or alternatively bonds and high-powered money) denominated in marks versus dollars. Frankel finds that the portfolio balance variables do not enter significantly in his quarterly equations, and indeed the key coefficients are of the wrong sign. In a subsequent paper Frankel (1982b) attempts to increase the power of his test by using monthly data, jointly estimating equations for six currencies (the U.S. Dollar, mark, pound sterling, yen, French franc, and the Canadian dollar and imposing theoretical restrictions based on a model in which investors in each country maximize a function of the mean and variance of wealth. Although the coefficients have the right sign, they remain jointly insignificant. Frankel's work follows that of Dooley and Isard (1983), who do not perform formal hypothesis tests but instead estimate a regression subject to a grid of prior constraints. They conclude that the exchange rate risk premium can account at most for only a small percentage of quarterly mark/dollar movements. Rogoff (1984) replicates Frankel's study, using high-frequency weekly data to detect a portfolio balance effect in the Canadian dollar/U.S. dollar exchange rate risk premium. The data have resisted all his efforts to obtain equations for the uncovered interest rate differential in which portfolio balance variables appear with statistically significant coefficients of the right sign. The results are therefore no more encouraging than those obtained in studies based on lower-frequency data. Loopesko (1984) tests the error orthogonality property of risk premium for a selection of currencies and finds that it does not hold for the majority of currencies. Loopesko uses a measure of cumulated daily data on official intervention for the G7 countries instead of a measure of outstanding asset stocks to estimate the portfolio balance model using the inverted demand approach. The joint hypothesis of rational expectations and perfect substitutability is tested by estimating an equation for the risk premium where the explanatory variables considered are lags of the dependent variable, lagged exchange rates and the cumulated intervention proxy variable. The coefficient on the lagged cumulated intervention variable is found to be statistically significantly different from zero at conventional nominal levels of significance for various sample periods considered. Loopesko concludes that the results provide evidence that sterilised intervention is short term effective through the portfolio balance channel.

By the early 1980s, some empirical successes in the literature had been overturned and key empirical findings began to turn negative. The most profound negative result was produced by Meese and Rogoff (1983), who compared the predictive abilities of a variety of exchange rate models. Their key result was that no existing structural exchange rate model could reliably predict the alternative of a random walk at short and medium run horizons, even when added by leaded values of the regressors. This extremely negative finding has never been entirely convincingly overturned despite many attempts. The simple random walk model of the exchange rate has become the standard benchmark for empirical exchange rate performance. Although the Meese and Rogoff finding is remarkably robust, a number of authors have found models whose out-of-sample forecasting performance improves upon a random walk (Mac Donald and Taylor (1993), Mark (1995) and Mac Donald and Marsh (1997)).

Obstfeld and Rogoff (1995), dismisses portfolio balance theory as partial equilibrium reasoning because it omits the government budget constraint. This point is made most comprehensively in an important paper by Backus and Kehoe (1989). Using only an arbitrage condition, they show that under complete asset markets, or under incomplete asset markets and a set of spanning conditions, changes in the currency composition of government debt require no offsetting changes in monetary and fiscal policies to satisfy both the government's and households' budget constraints.

A number of authors have attempted to detect an effect of sterilised intervention on the risk premium, using either the ratio of home to foreign outside assets [Frankel (1982,a, b), Rogoff (1984), Dankar et al. (1987)] or on actual intervention data [Loopesko (1984), Dominguez (1990)] as the explanatory variable. These papers expand on tests of foreign exchange market efficiency in that they try to provide a theoretical basis for the existence of the risk premium. On the whole, these studies suggest that unless daily official intervention figures are used the effects of sterilised intervention on the risk premium are either statistically insignificant or the coefficients have the wrong sign. Dominguez (1990) interprets these results as evidence of the signaling effects of intervention about future monetary policies as opposed to any portfolio balance effects.

According to the theory, as long as foreign and domestic assets are considered outside assets and are imperfect substitutes for each other in investor's portfolios, an intervention that changes the relative outstanding supply of domestic assets will require a change in expected relative returns. This is likely to result in a change in the exchange rate. Existing empirical evidence on the effectiveness of intervention is mixed: studies using data from the 1970s suggest that intervention operations that do not affect the monetary base have, at most, a short-lived influence on exchange rates, but more recent studies indicate that the intervention operations that followed the Plaza Agreement influenced both the level and variance of exchange rates. In the 1980s and early 1990s, attention focused on the effect of sterilized intervention on the level of the exchange rate and on the channels through which it works. The results on the effectiveness of intervention are mixed and depend on which exchange rate is analyzed, what sample period is studied and the intervention strategy that was used. In an influential paper, Dominguez and Frankel (1993) measure the risk premium using survey data and show that the resulting measure can in fact be explained by an empirical model which is consistent with the portfolio balance model with the additional assumption of mean-variance optimisation on the part of investors. The authors describe the econometric problems that arise in the standard portfolio balance estimation equation by deriving an alternative portfolio balance specification that measures the expected change in exchange rates using survey data rather than ex post exchange rate changes. The effectiveness of sterilised intervention is established both through the portfolio balance channel and the signaling channel. However, as long as the interest differential does not fully absorb the impact of intervention on the risk premium, their coefficient estimates indicate that foreign exchange interventions do matter. Thus, they provide evidence that official announcements of exchange rate policy and reported intervention significantly affect exchange rate expectations, so they implicitly support the portfolio balance model. Lewis (1995) examines the relationship between foreign exchange market intervention and monetary policy, testing the hypothesis that official intervention signals changes in future monetary policy as well as the hypothesis that changes in monetary policy may induce leaning-against-the-wind interventions. Lewis' study provides persuasive supportive evidence for both hypotheses, suggesting that official intervention may predict monetary policy variables and vice versa. However, other papers do not support the conclusion that intervention is effective. Humpage (1988), for example, concludes that intervention was unable to influence the dollar's level. Baillie and Osterberg (1997) find that over the period August 1985 to March 1990, Federal Reserve intervention did not influence the mark/dollar or yen/dollar exchange rates.

The empirical literature on exchange rates has highlighted a variable called currency order flow as strongly correlated with exchange rate returns (Evans and Lyons (2002), (2003a), Hau et al. (2002), Killeen et al. (2002), Rime (2001)). Order flow is sometimes interpreted as the variable through which dispersed information is aggregated and reflected in the price (Lyons (2001), Evans and Lyons (2003b)). Yet simple portfolio shifts could also give rise to order flow without any role for information asymmetries. Within the portfolio rebalancing framework and conditional on exogenous equity return and exchange rate shocks, it is plausible that net capital flows and order flows are closely aligned. Conditional on an exogenous appreciation of his foreign wealth for example, the home investor is likely to initiate the selling of foreign assets as well as the selling of foreign currency balances. According to Evans and Lyons (2002) portfolio model, trade innovations affect exchange rates through a portfolio-balance effect, given that foreign dealers are willing to absorb an excess demand or supply of foreign currency from their customers only if compensated by a shift in the exchange rate. While Evans and Lyons propose a formal model for their portfolio-shift effect, in their empirical investigation they do not directly test it. Instead, they estimate a reduced form specification. They prove that portfolio shocks in foreign exchange markets might also have long term effects on exchange rates.

Cushman (2006) applies the portfolio balance model to the Canadian–U.S. exchange rate over the floating period. Cointegration vectors that closely match theoretical expectations for the two countries' demands for home and foreign assets are found, and the exchange rate is important in the error correction process. The model is able to beat a random walk at some out-of-sample forecast horizons.

Financial globalization is stimulating researchers to take a fresh look at exchange rate economics. Lane and Milesi-Ferretti (2006) work out the implications of financial globalization for exchange rate behaviour. They conclude that a wider dispersion in net foreign asset positions implies stronger long-term trends in real exchange rates and that the impact of currency movements on net external wealth is an increasing function of the scale of international balance sheets.

Breedon and Vitale (2010) indicate that the strong contemporaneous correlation between order flow and exchange rates is largely due to portfolio-balance effects. They distinguish the information and portfolio-balance effects of order flow based on the direct estimation of a structural model of exchange rate determination, where trade innovations affect exchange rates via both their information content and their impact on the inventories of foreign investors. Their results indicate that the strong contemporaneous correlation between order flow and exchange rates is largely due to portfolio-balance effects.

Some researchers have attempted to improve the performance of both the monetary approach and the portfolio balance approach to exchange rate determination by considering empirical models which incorporate features of both approaches. Omitting risk in the monetary model, where no allowance is made for imperfect substitutability of non-monetary assets, may be an important source of misspecification. Since the exchange rate determined in the portfolio balance model is expected to balance the current account, agents forming rational expectations will revise their expected exchange rate when news on the future path of the current account become available. The hybrid monetary-portfolio balance model represents a synthesis of the features of the monetary model and the portfolio balance model in which allowance is made for a risk premium and news about the current account. Hooper and Morton (1982) assume that the risk premium depends upon the cumulated current account surplus net of the cumulation of foreign exchange market intervention. Their results, obtained using data for the dollar effective exchange rate in the 1970s, are mixed: the monetary model variables are statistically significant at conventional nominal levels of significance and correctly signed, but among the portfolio balance variables only the news term is statistically significant while the risk premium is statistically insignificant and wrongly signed. Nevertheless, Hooper and Morton interpret their results as an important improvement relative to monetary models and to portfolio models. Frankel (1984) estimates the hybrid monetary-portfolio balance model not including the news term and solving the model for the risk premium. He considers a model in which only two assets are held in the portfolio: those denominated in domestic currency, and those denominated in foreign currency (dollars). Frankel finds the risk premium to be statistically significant at conventional nominal levels of significance, while the monetary terms are often found to be not statistically significantly different from zero. Pilbeam (1995) found that there is little difference in the predictive success of the alternative exchange rate models, however, there are significant differences in the performance of a model depending upon the expectations mechanism specified. He proves that the flexible price monetary model, the portfolio balance model and a hybrid model under extrapolative and adaptive expectations mechanisms provide statistically significant information about the direction of exchange rate movements. By contrast, the same models when employing static, regressive and rational expectation mechanisms do not provide any statistically significant information.

3. The portfolio balance model

A distinguishing feature of the portfolio balance model among exchange rate models is the assumption of imperfect substitutability between domestic and foreign assets. The assumptions, present both in the flexible-price and the sticky-price monetary models, that domestic and foreign assets are perfect substitutes and that the wealth effects of current account imbalances are negligible are relaxed in the portfolio balance model (Dornbusch and Fischer, 1980; Isard, 1980; Branson, 1983, 1984).The level of the exchange rate is determined in the portfolio balance model by supply and demand for domestic and foreign financial assets. The exchange rate is the main determinant of the current account balance. That is to say, a surplus in the current account balance is associated with a rise in net domestic holdings of foreign assets, which influences the level of wealth, and in turn the level of the demand for assets, which ultimately affects the exchange rate. The portfolio balance model may be seen, therefore, as a dynamic model of exchange rate determination based on the interaction of asset markets, current account balance, prices and the rate of asset accumulation, which allows one to distinguish between the short-run and the long-run equilibrium. A large number of studies focus on the traditional formulation of the portfolio balance model under the assumption that the Ricardian equivalence theorem does not hold. In that framework, investors allocate their wealth among different assets in proportions that are assumed to be increasing functions of the expected return on each asset. Moreover, under the assumption that investors are risk-averse and that rates of return are uncertain, investors maximise expected profits by diversifying their portfolios.

The model economy we study is based on Branson and Henderson (1985) who provide in their paper a complete survey of portfolio-based approaches.

3.1 Short run equilibrium

The short run adjustment implies that we assume that both domestic prices and output are fixed. The model economy we study is a small open economy such that the rest of the world can be taken as given. The net financial wealth of the private sector (*W*) is divided into three components: money (*M*), domestically issued bonds (*B*) denominated in domestic currency and foreign bonds denominated in foreign currency and held by domestic residents (*B**). Using the exchange rate (S) we denominate the foreign bonds in domestic currency. Domestic bonds can be considered as government debt held by the domestic private sector and foreign bonds can be thought as the level of net claims on foreigners held by the private sector.

$$W \equiv M + B + SB^*$$

With domestic and foreign interest rates given by i and i^* we can express a definition of wealth and domestic demand functions.

$$M = M(i, i^* + E(s) - s)W$$
$$B = B(i, i^* + E(s) - s)W$$
$$sB^* = B^*(i, i^* + E(s) - s)W$$

The scale variable is the level of wealth and the demand functions are homogeneous in wealth. This allows them to be written in nominal terms (assuming homogeneity in prices and real wealth, prices cancel out). Goods prices are indeterminate in this model (in what follows we assume long-run neutrality).Since, under a free float, a current account surplus must be exactly matched by a capital account deficit the current account must determine the rate of accumulation of *B**over time :

$$\dot{B}^* = T(S/P) + i^* B^*$$

This equation gives the rate of change of *B**,the capital account, as equal to the current account, which is in turn equal to the sum of the trade balance, $T(\cdot)$ and net debt service receipts, i *B*. The trade balance depends positively on the level of the real exchange rate (a devaluation improves the trade balance).

An increase in money induces, ceteris paribus, agents to adjust their portfolios by buying both domestic and foreign bonds and a new equilibrium is established at a lower level of the interest rate and a higher level of the nominal exchange rate. An increase in domestic bonds causes interest rates to rise as the excess supply of domestic bonds tends to depress their market price. An increase in foreign bonds generates excess supply of foreign currency as agents try to adjust their portfolios. If domestic and foreign bonds are close substitutes the wealth effect may be expected to be swamped by the substitution effect of increased holdings of domestic bonds, and net sales of foreign bonds follow, ultimately inducing an appreciation of the domestic currency. Conversely, if domestic and foreign bonds are not close substitutes in portfolios, then the wealth effect dominates the substitution effect, leading to a depreciation of the domestic currency,as agents increase their holdings of foreign bonds.

It is clear that the change in the exchange rate caused by openmarket operations involving domestic assets is smaller relatively to open-market operations involving foreign assets, while the effect on the domestic interest rate is greater. The direct effect of the purchase of domestic assets is on the domestic interest rate, whereas the purchase of foreign assets affects the exchange rate more directly. Therefore, the impact of open-market operations on the exchange rate and interest rate may strongly depend on the combination of domestic and foreign assets purchased by the government in its open-market operations.

3.2 Long run equilibrium

In order to characterise long-run equilibrium, we need to make sure that there is no tendency for changes in the levels of the stocks of the various assets held by domestic residents. An increase in the money supply is expected to lead to a price increase, which will affect net exports and hence influence the current account balance. In turn, this will affect the level of wealth which feeds back into the asset market, affecting the exchange rate during the adjustment to long-run equilibrium. Assuming that the foreign price level is constant, the current account of the balance of payments, *CA*, expressed in foreign currency, is:

$$CA = T(S/P) + i^* B^*$$

Where the trade balance T (·) is a function of competitiveness (it improves if the exchange rate rises or the domestic price level falls. If the economy considered is a capital exporter and i*B* is positive, then a balance on the current account requires a deficit on the trade balance. Now consider the case by which the government purchases domestic bonds by printing money. In order to induce agents to hold more money and fewer bonds, the domestic interest rate falls and, as agents attempt to compensate for the reduction in their portfolios of the level of domestic assets by buying foreign bonds, the domestic currency will depreciate. The net impact effect is a lower domestic interest rate and a depreciated currency. Monetary policy, for example an open market operation leading to an increase in the money supply, has an immediate effect on the exchange rate and the interest rate. The exchange rate depreciates and the interest rate falls. Since the current account is the sum of the trade balance and net foreign investment income, a current account surplus is reflected by an accumulation of foreign bonds which feeds back into the asset markets leading to an appreciation of the currency. This is an additional effect on the exchange rate in the long run. This together with adjustments of the price level will affect the trade balance. In the long run, the current account is balanced such that a positive net foreign investment income requires a deficit in the trade balance.

4. Data & model

Based on the Branson and Henderson specification at the previous section -in the short run and in the long run- the models used here are the following linearized versions :

$$s_t = a_0 + a_1 b_t + a_2 b^{usa}{}_t + a_3 (i_t - i_t{*}) + u_t$$

$$s_t = \beta_0 + \beta_1 (b_t - b^{usa}{}_t) + \beta_2 (i_t - i_t{*}) + u_t$$

$$s_t = \gamma_0 + \gamma_1 CA_t + \gamma_2 CA^{usa}{}_t + \gamma_3 (i_t - i_t{*}) + v_t$$

$$s_t = \delta_0 + \delta_1 (CA_t - CA^{usa}{}_t) + \delta_2 (i_t - i_t{*}) + v_t$$

where s is the log value of the exchange rate, b and b^{usa} are the log values of the debt for the domestic and foreign country respectively, CA and CA^{usa} are the current account balances for the domestic and foreign country respectively, and i-i* denotes the rate difference between the domestic and foreign country. The analysis is specified for the Yen/US dollar and the Euro/US dollar exchange rate. Finally, static expectations are assumed, thus the term E(s) is zero. Complete definitions and data sources are given in appendix.

5. Unit root tests

Non stationarity of the economic variables involved in the analysis leads to violation of the classical assumptions of standard regression methods, and to spurious estimates. The possible endogeneity of regressors is a problem not well handled by OLS. The sample sizes available for data analysis are usually small leading to small sample bias in estimates. Classical regression properties hold only for cases where variables are stationary (integrated of order 0), that by contrast most economic variables are integrated of order 1 or higher (and hence do not satisfy these assumptions), but that where error correction mechanisms or long run relationships exists, certain combinations of I(1) variables are likely to be I(0) and hence amenable to OLS estimation. Where this is so, the variables are said to be cointegrated and OLS estimates of such cointegrated variables may be superconsistent in the sense of collapsing to their true values more quickly than if the variables had been stationary. The first step is to determine the degree of integration of the individual series under investigation, thus the empirical analysis begins by examining this with univariate tests. Where a cointegrating relationship cannot be found, no long run relationship among the variables can be demonstrated and we have the case of spurious regression

A variable is determined to be I(1) if a unit root is found in levels and stationarity is found in first differences. The tests employed are for both the unit root null and the stationarity null. Most of the variables in levels appear to possess trends of some sort, and so a linear trend is included in the unit root tests on levels.

5.1 Dickey-Fuller (ADF) test

The early work on testing for a unit root in time series was done by Dickey and Fuller (Fuller, 1976; Dickey and Fuller, 1979).The basic objective of the test is to examine the null hypothesis that $\varphi = 1$

in $y_t = \phi y_{t-1} + u_t$

against the one-sided alternative $\varphi < 1$.

Thus the hypotheses of interest are:

H_0: series contains a unit root

H_1: series is stationary

Dickey-Fuller (DF) tests are also known as t -tests, and can be conducted allowing for an intercept, or an intercept and deterministic trend, or neither, in the test regression. The test statistics do not follow the usual t-distribution under the null hypothesis, since the null is one of non-stationarity, but rather they follow a non-standard distribution. The null hypothesis of a unit root is rejected in favour of the stationary alternative in each case if the test statistic is more negative than the critical value. The tests above are valid only if u_t is white noise. In particular, u_t is assumed not to be autocorrelated, but would be so if there was autocorrelation in the dependent variable of the regression (Δy_t) which has not been modelled. If this is the case, the test would be oversized, meaning that the true size of the test would be higher than the nominal size used. The solution is to augment the test using p lags of the dependent variable. The alternative model in the previous case is now written :

$$\Delta y_t = \psi \, y_{t-1} + \sum_{i=1}^{p} a_i \Delta y_{t-i} + u_t$$

The lags of Δy_t now soak up any dynamic structure present in the dependent variable, to ensure that u_t is not autocorrelated. The test is known as an augmented Dickey-Fuller (ADF) test and the same critical values from the DF tables are used as before.

5.2 ADF-GLS test

Elliott, Rothenberg, and Stock (1996) introduce a potentially more powerful unit root test: the generalized least squares version of the ADF test. They find that powers of ADF tests are lower than those of the limiting power functions when deterministic components (mean or trend) are included in the data generating process.

5.3 Lag selection procedure

A problem now arises in determining the optimal number of lags of the dependent variable. Although several ways of choosing *p* have been proposed, they are all somewhat arbitrary, and are thus not used here. Instead, an information criterion can be used to decide the number of lags. Including too few lags will not remove all of the autocorrelation, thus biasing the results, while using too many will increase the coefficient standard errors. The latter effect arises since an increase in the number of parameters to estimate uses up the degrees of freedom. Therefore, everything else being equal, the absolute values of the test statistics will be reduced. This will result in a reduction in the power of the test, implying that for a stationary process the null hypothesis of a unit root will be rejected less frequently than would otherwise have been the case. Ng and Perron (1995) suggest the lag length selection procedure that results in stable size of the test and minimal power loss. First, set an upper bound p_{max} for p and then estimate the ADF test regression with $p = p_{max}$. Next, reduce the lag length by one and repeat the process. For the choice of the lags, it is used the general-to-specific method by Ng and Perron and information criteria, i.e., the Akaike information criterion (AIC) and the Bayesian information criterion (BIC).

5.4 KPSS test

The most important criticism that has been levelled at unit root tests is that their power is low if the process is stationary but with a root close to the non-stationary boundary especially with small sample sizes. One way to get around this problem is to use a stationarity test as well as a unit root test. One such stationarity test is the KPSS test (Kwaitkowski et al., 1992). The asymptotic distribution of the statistic is derived under the null and under the alternative that the series is difference-stationary. The results of these tests can be compared with the ADF procedure to see if the same conclusion is obtained. For the conclusions to be robust both tests concluded that the series is stationary or non-stationary, respectively. By testing both the unit root hypothesis and the stationarity hypothesis, we can distinguish series that appear to be stationary, series that appear to have a unit root, and series for which the data (or the tests) are not sufficiently informative to be sure whether they are stationary or integrated.

If the hypothesis of nonstationarity of the individual series is rejected, then we cannot go any further. By contrast, if the hypothesis is not rejected, then it is correct and advisable to test for a unit root on the first difference of the series in question in order to exactly specify the order of integration. Only at this stage might the possibility of cointegration arise. If a linear combination between two or more series is defined which reduces the order of integration, the variables in the estimated regression are said to be cointegrated. Typically, we are dealing with I(1) series and I(0) linear combinations. Nevertheless, if the no-cointegration hypothesis cannot be rejected, then the estimated regression is spurious and has no economic meaning.

5.5 Unit root tests results

Here, the unit root null is examined with two tests: (1) the ADF test and (2) the DF-GLS test of Elliott, Rothenberg, and Stock (1996). The stationarity null is tested with the KPSS test (Kwiatkowski, et al., 1992). The results of unit root tests are presented in the tables below. In order to determine the lag order of the unit root tests the Ng and Perron procedure is applied starting with 8 lags by using the Bayesian information criterion. The optimal number of lags for each variable is presented in the second column of the following tables. Using the ADF test most of the variables, except the Japan-USA rate differential, the eurozone net debt and the eurozone-USA rate differential which are I(2) and the Yen/Dollar exchange rate which is stationary in level, are considered to be I(1). While, implementing the ADF-GLS test five of the variables (eurozone net debt, eurozone-USA rate differential, Japan-USA rate differential, Yen/Dollar exchange rate and Japan current account balance) are I(0) and the rest of them are I(1). Finally, testing the stationarity null all the variables are I(1) except the eurozone net debt, the Japan-USA rate differential and the USA private debt which are stationary in levels in 10% level.

6. Methodology & results

6.1 Methodology

With regard to the estimation of cointegrating regression models, it is well known that the ordinary least squares (OLS) estimator contains the second-order bias, including the endogeneity bias and the non-centrality bias, when the I(1) regressors are endogenous and the regression errors are serially correlated. Parameter estimates can be biased in small samples as well as in the presence of dynamic effects, and this bias varies inversely with the size of the sample and the calculated R^2. Also, when the number of regressors exceeds two there can be more than one cointegrating relationship or vector and it is difficult to give economic meaning to this finding. Then there is the problem caused by the likely endogeneity of the regressors, which would prevent OLS estimating the true values of the parameters. These difficulties associated with the OLS approach have led to the development of alternative procedures which are proposed in the literature.

One method of extracting purely the long-run coefficients is by the Johansen and Juselius procedure which is based on a maximum-likelihood approach. Another more recent and more robust method, (particularly in small samples) proposed by Stock and Watson (1993), which also corrects for possible simultaneity bias among the regressors, involves estimation of long-run equilibria via dynamic OLS (DOLS). Stock and Watson (1993) suggest a parametric approach for estimating long-run equilibria in systems which may involve variables integrated of different orders but still cointegrated. The potential of simultaneity bias and small-sample bias among the regressors is dealt with by the inclusion of lagged and led values of the change in the regressors. Their method improves on OLS by coping with small sample and dynamic sources of bias. The Johansen method, being a full information technique, is exposed to the problem that parameter estimates in one equation are affected by any mispecification in other equations. The Stock-Watson method is, by contrast, a robust single equation approach which corrects for regressor endogeneity by the inclusion of leads and lags of first differences of the regressors, and for serially correlated errors by a GLS procedure. In addition it has the same asymptotic optimality properties as the Johansen distribution. Furthermore, based on Monte Carlo evidence, Stock and Watson (1993) show that DOLS is more favourable, particularly in small samples, compared to a number of alternative estimators of long-run parameters, including those proposed by Engle and Granger (1987), Johansen (1988) and Phillips and Hansen (1990).

In estimating the parameters, the DOLS procedure is adopted where the log of the exchange rate is regressed on the levels of the independent variables plus the lags and leads of their first differences. The models that are used (discussed at the section 4, based on Branson and Henderson specification) are the following ones :

$$s_t = a_0 + a_1 b_t + a_2 b^{usa}{}_t + a_3(i_t - i_t{*}) + \sum_{j=-k}^{j=k} \psi_j \Delta b_{t-j} + \sum_{j=-k}^{j=k} \phi_j \Delta b^{usa}{}_{t-j} + \sum_{j=-k}^{j=k} \zeta_j \Delta(i_{t-j} - i_{t-j}{*}) \ + u_t$$

$$s_t = \beta_0 + \beta_1(b_t - b^{usa}{}_t) + \beta_2(i_t - i_t{*}) + \sum_{j=-k}^{j=k} \psi_j \Delta(b_{t-j} - b^{usa}{}_{t-j}) + \sum_{j=-k}^{j=k} \phi_j \Delta(i_{t-j} - i_{t-j}{*}) + u_t$$

$$s_t = \gamma_0 + \gamma_1 CA_t + \gamma_2 CA^{usa}{}_t + \gamma_3 (i_t - i_t*) + \sum_{j=-k}^{j=k} \psi_j \Delta CA_{t-j} + \sum_{j=-k}^{j=k} \phi_j \Delta CA^{usa}{}_{t-j} + \sum_{j=-k}^{j=k} \zeta_j \Delta (i_{t-j} - i_{t-j}*) + v_t$$

$$s_t = \delta_0 + \delta_1 (CA_t - CA^{usa}{}_t) + \delta_2 (i_t - i_t*) + \sum_{j=-k}^{j=k} \psi_j \Delta (CA_{t-j} - CA^{usa}{}_{t-j}) + \sum_{j=-k}^{j=k} \phi_j \Delta (i_{t-j} - i_{t-j}*) + v_t$$

Following the work of Hayakawa and Kurozumi (2008) the dynamic OLS estimator without leads substantially outperforms that with leads and lags when the cointegrating error does not Granger-cause the first difference of I(1) variables that appear in the right side. The authors investigate the case where leads are unnecessary for the DOLS method, and by using the Monte Carlo simulation they demonstrate that in such a case, we can expect the improvement of the DOLS estimator in terms of the mean squared error by excluding leads from the regressors. They also consider verifying the DOLS estimation without leads by investigating whether or not the regression error from the dynamic OLS regression without leads is serially uncorrelated. For this purpose, the portmanteau tests are available as it is explained in Lutkepohl (1993). Performing the portmanteau test (Ljung–Box autocorrelation test) no autocorrelation is found on the Eurozone/U.S.A. DOLS regressions. The P-value using the U.S. net government debt is 0.301 while using the U.S. private debt is 0.11, hence the null hypothesis of no autocorrelation is accepted for both regressions (additionally, the $X^2(1)$ statistic is 1.06787 and 2.09118 respectively). Thus, according to Hayakawa and Kurozumi the leads are unnecessary and the model becomes :

$$s_t = a_0 + a_1 b_t + a_2 b^{usa}{}_t + a_3 (i_t - i_t*) + \sum_{j=0}^{j=k} \psi_j \Delta b_{t-j} + \sum_{j=0}^{j=k} \phi_j \Delta b^{usa}{}_{t-j} + \sum_{j=0}^{j=k} \zeta_j \Delta (i_{t-j} - i_{t-j}*) \ + u_t$$

Nevertheless, performing the portmanteau test on the Japan/U.S.A. DOLS regressions autocorrelation is found (the null hypothesis of no autocorrelation is rejected). The P-values using the debt(net government and private), the debt difference, the current account balance and the current account difference models are 2.52e-08,4.37e-08,1.95e-08,4.34e-17 and 7.58e-20 respectively (additionally the $X^2(1)$statistics are 31.0466,29.9792,31.5458,70.6171 and 83.1555). Thus, the DOLS regressions include lags and leads of the first differences of the independent variables.

Starting with 3 lags and 3 leads for the Japan regressions and with 3 lags for the Eurozone regressions restrictions are set for the third lagged and led terms in every regression. The restrictions that the third lagged and led terms are zero are rejected, so all the DOLS regressions about Eurozone include 3 lags and all the DOLS regressions about Japan include 3 lags and 3 leads.

6.2 Results

In table 5 there are presented the results of regressing the Yen/U.S. dollar exchange rate on the net U.S. government debt, the net Japan government debt and on the interest rate difference (defined as the Japan interest rate minus the U.S. interest rate). The interest rate difference is statistical insignificant while both U.S. and Japan debt which are statistical significant have the wrong sign. In table 6 the Yen/U.S. dollar exchange rate is regressed on the private U.S. debt, the net Japan government debt and on the interest rate difference. All the variables are statistical significant but all the signs are wrong. In table 7 the exchange rate is regressed on the debt difference (defined as the Japan net debt minus the net government U.S. debt) and on the interest rate difference. Both the variables are statistical significant and have the expected signs. The statistical significance of the debt difference demonstrates a portfolio effect where the imperfect substitutability of assets is implied. The rise of the domestic interest rate leads to an increase in the demand for domestic assets, which causes the demand for foreign assets to decrease as agents adjust their portfolios. The decrease of the demand for foreign assets leads to a decrease of the demand for foreign currency and thus an appreciation of the domestic currency (fall of the exchange rate). In table 8 the Yen/U.S. dollar exchange rate is regressed on the current account balances of the two countries (a deficit for the U.S. and a surplus for the Japan) and on the rate difference. The variables are statistical significant but only the rate difference has the expected sign. The current account balance affects the level of wealth which feeds back into the asset market, affecting the exchange rate during the adjustment to long-run equilibrium. A current account surplus in the domestic country (here in Japan) is reflected by an accumulation of foreign bonds (through the portfolio rebalancing) which feeds back into the asset markets leading to an appreciation of the domestic currency (decrease of the exchange rate). A current account deficit in the foreign country (here in U.S.) leads to an accumulation of domestic bonds (through the portfolio rebalancing) which causes a depreciation of the foreign currency (U.S. dollar) and a decrease of the exchange rate. In table 9 the exchange rate is regressed on the current account difference (defined as the U.S. current account balance minus the Japan current account balance) and on the interest rate difference. Both the variables are statistical significant and have the correct signs. In table 10 the U.S. dollar/Euro exchange rate is regressed on the net U.S. government debt, the net eurozone government debt and on the interest rate difference (defined as the eurozone interest rate minus the U.S. interest rate) and all the variables are statistical significant and have the correct signs. In table 11 the U.S. dollar/Euro exchange rate is regressed on the private U.S. debt, the net eurozone government debt and on the interest rate difference where again all the variables are statistical significant and have the expected signs.

7. Conclusions

An empirical portfolio balance model based on Branson and Henderson is specified for the Japan/U.S. exchange rate and the Eurozone/U.S. exchange rate. Empirical implementation using dynamic OLS techniques reveals that for the Japan/U.S. exchange rate the portfolio balance effect works only when it is regressed on the debt and current account differences while the regressions for the Eurozone/U.S. exchange rate on the domestic and foreign debt have the expected results verifying the portfolio balance model in the short run period. Unfortunately, the results are not satisfactory when the Japan/U.S. exchange rate is regressed on the domestic and foreign debt and on the domestic and foreign current account balances.

8. Tables

Table 1: unit root test results for Japan (levels)

Variable	lags	ADF (P-value)	ADF-GLS (P-value)	KPSS
exchange rate	3	-3.49085 (0.04029)**	-2.77551 (0.00536)***	0.49761
rate difference	2	-2.63321 (0.2653)	-2.56496 (0.01)***	0,183191*
Japan debt	4	-2.95568 (0.145)	-0.24087 (0.5997)	0.23234
USA debt	1	-0.3107 (0.9904)	2.53243 (0.9976)	0.61421
USA private debt	6	-2.72725 (0.2254)	-1.1274 (0.2367)	0,511032*
debt difference	3	-1.6054 (0.7913)	0.33742 (0.7827)	1.26611
CAB Japan	2	-2.91362 (0.158)	-2.63411 (0.00818)***	0.44664
CAB USA	2	-2.00094 (0.6003)	-0.07902 (0.6564)	0.39021
CAB difference	3	-2.26586 (0.4522)	-0.94302 (0.3085)	0.46307

Significant results are highlighted with (*) for the 0.10 level, (**) for the 0.05 level, and (***) for the 0.01 level.

Table 2: unit root test results for Japan (first differences)

Variable	lags	ADF (P-value)	ADF-GLS (P-value)	KPSS
exchange rate	3	-3.18705 (0.08697)*	-2.96126 (0.0029)***	0,04982***
rate difference	2	-2.41242 (0.3729)	-2.4623 (0.0133)**	0,11148***
Japan debt	4	-3.23756 (0.07711)*	-2.94576 (0.0031)***	0,05159***
USA debt	1	-4.72828 (0.00058)***	-2.94892 (0.0031)***	0,11955***
USA private debt	6	-3.1339 (0.09835)*	-2.23772 (0.02434)**	0,22525***
debt difference	3	-6.62518 (0.00000)***	-2.847 (0.0043)***	0,07929***
CAB Japan	2	-5.429 (0.00002)***	-1.26462 (0.1902)	0,04630***
CAB USA	2	-6.36154 (0.00000)***	-4.45997 (0.0001)***	0,05767***
CAB difference	3	-3.59313 (0.03034)***	-1.75265 (0.07567)*	0,10297***

Significance is highlighted as in Table 1.

Table 3: unit root test results for Eurozone (levels)

Variable	lags	ADF (P-value)	ADF-GLS (P-value)	KPSS
exchange rate	0	-2,9375 (0,1628)	0,1147 (0,7189)	0,3717
rate difference	1	-3,0565 (0,117)	-2,8903 (0,0037)***	0,2775
eurozone debt	4	-2,7086 (0,233)	-1,6764 (0,0887)*	0,164332*
USA debt	1	-0,3107 (0,9904)	2,53243 (0,9976)	0,61421
USA private debt	6	-2,7272 (0,2254)	-1,1274 (0,2367)	0,511032*

Significance is highlighted as in Table 1.

Table 4: unit root test results for Eurozone (first differences)

Variable	lags	ADF (P-value)	ADF-GLS (P-value)	KPSS
exchange rate	0	-4,9292 (0,0016)***	-3,4722 (0,0005)***	0,190171*
rate difference	1	-2,7238 (0,2267)	-2,4206 (0,015)**	0,134948**
Eurozone debt	4	-1,613 (0,7882)	-1,862 (0,0597)*	0,0950297***
USA debt	1	-4,7282 (0,00058)***	-2,94892 (0,00311)***	0,119557***
USA private debt	6	-3,1339 (0,09835)*	-2,23772 (0,02434)**	0,225253***

Significance is highlighted as in Table 1.

Table 5 : Stock-Watson dynamic OLS parameter estimates of Yen/U.S. dollar exchange rate in the short run period.

Variable (expected sign)	Estimated coefficient (SE)
Japan debt $_t$ (+)	-0,344775 *** (0,0296006)
USA debt $_t$ (-)	1,03109 *** (0,0402217)
rate difference $_t$ (-)	0,0219399 (0,0131154)
Δ Japan debt $_t$	0,0932266 (0,076986)
Δ USA debt $_t$	-0,93292 (0,693676)
Δ rate difference $_t$	-0,00196737 (0,0167178)
Δ Japan debt $_{t+1}$	-0,169113 ** (0,0686966)
Δ Japan debt $_{t+2}$	-0,273053 *** (0,0655718)
Δ Japan debt $_{t+3}$	-0,336594 *** (0,085795)
Δ Japan debt $_{t-1}$	0,170247 ** (0,081311)
Δ Japan debt $_{t-2}$	0,0433359 (0,0530509)
Δ Japan debt $_{t-3}$	-0,0846339 *** (0,0276377)
Δ USA debt $_{t+1}$	2,53651 ** (0,940219)
Δ USA debt $_{t+2}$	2,07309 * (1,10128)
Δ USA debt $_{t+3}$	-1,61989 * (0,819881)
Δ USA debt $_{t-1}$	0,778682 (0,659692)
Δ USA debt $_{t-2}$	-3,50769 *** (1,16367)
Δ USA debt $_{t-3}$	-4,00778 *** (1,08304)
Δ rate difference $_{t+1}$	0,00600452 (0,0170457)
Δ rate difference $_{t+2}$	-0,0166262 (0,0165383)

Δ rate difference $_{t+3}$	0,0277953
	(0,0174308)
Δ rate difference $_{t-1}$	-0,0539017 ***
	(0,0173403)
Δ rate difference $_{t-2}$	0,0376604 *
	(0,0185825)
Δ rate difference $_{t-3}$	0,00699246
	(0,0236404)
Sum of squared residuals	0,0582169
Adjusted R^2	0,99987

Significance is highlighted as in Table 1.

Table 6 : Stock-Watson dynamic OLS parameter estimates of Yen/U.S. dollar exchange rate in the short run period (using the U.S. Private debt)

Variable (expected sign)	Estimated coefficient (SE)
Japan debt $_t$ (+)	-0,140381 *** (0,0255041)
USA private debt $_t$ (-)	0,824877 *** (0,0380876)
rate difference $_t$ (-)	0,0699454 *** (0,0136916)
Δ Japan debt $_t$	-0,119314 (0,0902545)
Δ USA debt $_t$	-0,916663 *** (0,317942)
Δ rate difference $_t$	-0,0101953 (0,0201637)
Δ Japan debt $_{t+1}$	-0,184076 ** (0,0769078)
Δ Japan debt $_{t+2}$	-0,284227 *** (0,0735487)
Δ Japan debt $_{t+3}$	-0,391319 *** (0,0998649)
Δ Japan debt $_{t-1}$	0,000464363 (0,0851583)
Δ Japan debt $_{t-2}$	-0,0725923 (0,0640625)
Δ Japan debt $_{t-3}$	-0,114235 *** (0,031783)
Δ USA private debt $_{t+1}$	0,744525 (0,593143)
Δ USA private debt $_{t+2}$	-0,413476 (0,505826)
Δ USA private debt $_{t+3}$	-1,93134 *** (0,572221)
Δ USA private debt $_{t-1}$	1,38554 ** (0,599043)
Δ USA private debt $_{t-2}$	-0,626782 (0,516814)
Δ USA private debt $_{t-3}$	-1,67988 ** (0,652434)
Δ rate difference $_{t+1}$	0,0455037 *** (0,0157143)
Δ rate difference $_{t+2}$	0,0385378 ** (0,0153428)

Δ rate difference $_{t+3}$	0,0548434 ***
	(0,0182387)
Δ rate difference $_{t-1}$	-0,0802138 ***
	(0,0150125)
Δ rate difference $_{t-2}$	0,00902426
	(0,0172969)
Δ rate difference $_{t-3}$	0,013966
	(0,0227665)
Sum of squared residuals	0,0685441
Adjusted R^2	0,99984

Significance is highlighted as in Table 1.

Table 7 : Stock-Watson dynamic OLS parameter estimates of Yen/U.S. dollar exchange rate in the short run period (using the debt difference)

Variable (expected sign)	Estimated coefficient (SE)
debt difference $_t$ (+)	0,366785 *** (0,0112035)
rate difference $_t$ (-)	-0,102268 *** (0,030746)
Δ debt difference $_t$	0,708673 *** (0,145201)
Δ rate difference $_t$	0,106501 ** (0,0467644)
Δ debt difference $_{t+1}$	0,993467 *** (0,12682)
Δ debt difference $_{t+2}$	0,922617 *** (0,134772)
Δ debt difference $_{t+3}$	1,06256 *** (0,244067)
Δ debt difference $_{t-1}$	0,430703 *** (0,14331)
Δ debt difference $_{t-2}$	0,26315 * (0,146739)
Δ debt difference $_{t-3}$	-0,0184869 (0,118751)
Δ rate difference $_{t+1}$	-0,0515816 (0,0788962)
Δ rate difference $_{t+2}$	-0,0515506 (0,0538764)
Δ rate difference $_{t+3}$	-0,0623942 (0,0419897)
Δ rate difference $_{t-1}$	0,260815 *** (0,0752538)
Δ rate difference $_{t-2}$	0,247208 *** (0,0530234)
Δ rate difference $_{t-3}$	0,201209 *** (0,0283097)
Sum of squared residuals	1,02583
Adjusted R^2	0,99836

Significance is highlighted as in Table 1.

Table 8 : Stock-Watson dynamic OLS parameter estimates of Yen/U.S. dollar exchange rate in the long run period

Variable (expected sign)	Estimated coefficient (SE)
Japan CAB $_t$ (-)	1,98811 ***
	(0,0815827)
USA CAB $_t$ (-)	0,574924 ***
	(0,08842)
rate difference $_t$ (-)	-0,236558 ***
	(0,0666113)
Δ Japan CAB $_t$	-1,53572 ***
	(0,281464)
Δ USA CAB $_t$	-1,15729 ***
	(0,327849)
Δ rate difference $_t$	-0,0181792
	(0,171704)
Δ Japan CAB $_{t+1}$	0,697159 ***
	(0,225161)
Δ Japan CAB $_{t+2}$	1,19793 ***
	(0,292329)
Δ Japan CAB $_{t+3}$	1,40149 ***
	(0,223288)
Δ Japan CAB $_{t-1}$	-1,36724 ***
	(0,255068)
Δ Japan CAB $_{t-2}$	-1,31955 ***
	(0,240796)
Δ Japan CAB $_{t-3}$	-1,23711 ***
	(0,251467)
Δ USA CAB $_{t+1}$	-0,762503 ***
	(0,274131)
Δ USA CAB $_{t+2}$	-0,516731 *
	(0,273523)
Δ USA CAB $_{t+3}$	-0,287682
	(0,362792)
Δ USA CAB $_{t-1}$	-0,521558 *
	(0,262787)
Δ USA CAB $_{t-2}$	0,190724
	(0,201702)
Δ USA CAB $_{t-3}$	0,178739
	(0,233677)
Δ rate difference $_{t+1}$	0,0602052
	(0,180499)
Δ rate difference $_{t+2}$	0,428586 *
	(0,244799)

Δ rate difference $_{t+3}$	0,439909 **
	(0,206029)
Δ rate difference $_{t-1}$	-0,0539018
	(0,184255)
Δ rate difference $_{t-2}$	0,466957 **
	(0,202079)
Δ rate difference $_{t-3}$	0,474764 **
	(0,215697)
Sum of squared residuals	41,4322
Adjusted R^2	0,97221

Significance is highlighted as in Table 1.

Table 9 : Stock-Watson dynamic OLS parameter estimates of Yen/U.S. dollar exchange rate in the long run period (using the current account difference)

Variable (expected sign)	Estimated coefficient (SE)
CAB difference $_t$ (-)	-0,474647 ***
	(0,103782)
rate difference $_t$ (-)	-0,518514 ***
	(0,162652)
Δ CAB difference $_t$	0,498807 *
	(0,285218)
Δ rate difference $_t$	0,431605
	(0,309206)
Δ CAB difference $_{t+1}$	-0,11041
	(0,313691)
Δ CAB difference $_{t+2}$	-0,106537
	(0,338413)
Δ CAB difference $_{t+3}$	-0,381513
	(0,306212)
Δ CAB difference $_{t-1}$	0,466876
	(0,298472)
Δ CAB difference $_{t-2}$	0,490729
	(0,343616)
Δ CAB difference $_{t-3}$	0,25185
	(0,312355)
Δ rate difference $_{t+1}$	-0,547496 *
	(0,28247)
Δ rate difference $_{t+2}$	-0,338266
	(0,322561)
Δ rate difference $_{t+3}$	-0,593116 *
	(0,328668)
Δ rate difference $_{t-1}$	0,523517 *
	(0,270469)
Δ rate difference $_{t-2}$	0,56082 *
	(0,307657)
Δ rate difference $_{t-3}$	0,792883 **
	(0,370659)
Sum of squared residuals	194,6
Adjusted R^2	0,88398

Significance is highlighted as in Table 1.

Table 10 : Stock-Watson dynamic OLS parameter estimates of U.S. dollar/Euro exchange rate in the short run period

Variable (expected sign)	Estimated coefficient (SE)
Eurozone debt $_t$ (-)	-0,88044 ***
	(0,0940723)
USA debt $_t$ (+)	0,841372 ***
	(0,093784)
rate difference $_t$ (+)	0,0660871 ***
	(0,0197096)
Δ Eurozone debt $_t$	2,06726 ***
	(0,397749)
Δ USA debt $_t$	-3,48649 **
	(1,36535)
Δ rate difference $_t$	-0,0680901 *
	(0,0358741)
Δ Eurozone debt $_{t-1}$	1,62882 ***
	(0,384448)
Δ Eurozone debt $_{t-2}$	1,4714 ***
	(0,382274)
Δ Eurozone debt $_{t-3}$	1,44771 ***
	(0,397403)
Δ USA debt $_{t-1}$	-4,48204 ***
	(1,50817)
Δ USA debt $_{t-2}$	-3,69735 **
	(1,65094)
Δ USA debt $_{t-3}$	-1,04543
	(1,14501)
Δ rate difference $_{t-1}$	-0,103291 *
	(0,0553366)
Δ rate difference $_{t-2}$	-0,0601804
	(0,0556073)
Δ rate difference $_{t-3}$	-0,16318 **
	(0,0567748)
Sum of squared residuals	0,0276341
Adjusted R^2	0,96897

Significance is highlighted as in Table 1.

Table 11 : Stock-Watson dynamic OLS parameter estimates of U.S. dollar/Euro exchange rate in the short run period (using the U.S. Private debt)

Variable (expected sign)	Estimated coefficient (SE)
Eurozone debt $_t$ (-)	-1.01885 *** (0.223719)
USA private debt $_t$ (+)	1.01005 *** (0.222703)
rate difference $_t$ (+)	0.0902007 * (0.0440507)
Δ Eurozone debt $_t$	1.96904 *** (0.600099)
Δ USA private debt $_t$	-1.91613 * (1.06144)
Δ rate difference $_t$	-0.0634073 (0.0536846)
Δ Eurozone debt $_{t-1}$	1.51478 ** (0.658808)
Δ Eurozone debt $_{t-2}$	1.32009 ** (0.511245)
Δ Eurozone debt $_{t-3}$	1.05809 ** (0.419587)
Δ USA private debt $_{t-1}$	-1.71209 (1.10349)
Δ USA private debt $_{t-2}$	-1.90404 (1.4144)
Δ USA private debt $_{t-3}$	-0.83447 (1.11946)
Δ rate difference $_{t-1}$	-0.106986 (0.0616176)
Δ rate difference $_{t-2}$	-0.0460542 (0.0539367)
Δ rate difference $_{t-3}$	-0.158751 * (0.0841166)
Sum of squared residuals	0.0318759
Adjusted R^2	0.964202

Significance is highlighted as in Table 1

Appendix : Data definitions and sources

U.S. net debt is defined as “Federal government debt: total public debt” (Federal Reserve Bank of St. Louis, FRED) minus the monetary base M1 (IMF’s International Financial Statistics) denominated in foreign currency (U.S. dolars). U.S. private debt is defined as “federal debt held by private investors” (Federal Reserve Bank of St. Louis, FRED) denominated in foreign currency (U.S. dolars). Eurozone's net debt is defined as “gross debt at face value” minus the monetary base M1 (IMF’s International Financial Statistics) denominated in domestic currency (euros). Japan's net debt is calculated as “government bonds” (Japan ministry of finance) minus the monetary base M1 (IMF’s International Financial Statistics) denominated in domestic currency (Yen). The current account -deficit for USA and surplus for Japan- is defined as a percentage of GDP (OECD). The interest rates for U.S.A., Japan and Eurozone are the discount rates (IMF’s International Financial Statistics). Finally, the exchange rates are taken from the IMF’s International Financial Statistics.

The frequency of data is quarterly and the range from 2000 Q1 to 2008 Q3 for the Eurozone and from 1996 Q2 to 2008 Q3 for the short run analysis and from 1985 Q1 to 2008 Q3 for the long run analysis of Japan.

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