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Current Account Sustainability in Seven Developed Countries

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Abstract. This paper is an attempt to examine the G-7 sustainability properties of current accounts of seven developed countries, using a methodology based on fractional processes. The purpose of this study is to test for the sustainability of current account deficits in seven developed countries for the 1974:1-2001:3 period. The results indicate that all countries' current account is covariance non-stationary and three countries' (France, Italy and Canada) current accounts are mean reverting so that they are sustainable in the long run, while those of Germany, UK, US and Japan are not mean reverting and are unsustainable. These results should also signal a warning to creditors and policymakers, unless there are policy distortions or permanent productivity shocks to the domestic economies. Furthermore, persistent deficits may lead to increased domestic interest rates to attract foreign capital and, in addition to this, the accumulation of external debt owing to persistent deficits will imply increasing interest payments that impose an excess burden on future generations.

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1. Introduction

The current account for individual countries is a barometer for both policymakers and investors as it represents an indicator of economic performance (Baharumshah et al. 2003, Goldberg, et al. 1995). Policymakers and investors are interested in the aggravation of current accounts; and they believe that the current build up of claims on the country by foreigners violates the solvency condition with respect to the rest of the world. However, economists are more interested in the country's intertemporal solvency constraint than in indicating the size of current account deficits at any particular point in time. This constraint emphasizes the long-run path of the current account (Husted, 1992; Wu et. al., 1996). Recently, IMF (2002) warned that the current account deficit in the US, at current or higher rates, would take the US net foreign liability position to over-increasing levels, and that at same point an adjustment will be needed. Leachman and Francis (2000) argue that more recently the US may be running a Ponzi gamble against global capital markets. Such a gamble can be welfare improving, even in dynamically efficient economies, as long as the rate of growth remains above the real interest rate on external debt. However, should the gamble fail, some generations will be less well off and sustainability of current account may be an issue.

Temporary current account deficits, reflecting reallocation of capital to the country in which capital is more productive, are not as bad as they are thought to be. On the other hand, persistent deficits may have some important effects. Firstly, foreign capital inflows may increase the domestic interest rates. Secondly, the accumulation of external debt will probably reflect the increasing interest payments, which might mean an excess burden in the future (Hakkio, 1995; Wu, 2000). Moreover, persistent current account deficits might serve as a leading indicator of financial crises (Baharumshah et al. 2003).

The series of economic crises in the 1990s showed that large and persistent current account deficits can generate a favourable environment for external crises, especially when those deficits are financed through short-term capital inflows. In order to evaluate the stable long-run trajectory of a country's current account, two inter-related questions are needed to be addressed. The first one concerns the solvency properties of the debtor country; the second one concerns the sustainability properties of the current account deficit (North, 2002). Intertemporal solvency investigates the country's ability to repay its external debt. For a country to become intertemporally solvent, the present discounted value of future trade surpluses must be equal to the present value of its foreign debt (Milesi-Ferretti and Razin, 1996). This is the concept of intertemporal solvency implying that all debts will be repaid in the long-run. This issue has been quite extensively covered in recent empirical research by implementation of applied stationarity and cointegration tests to examine intertemporal solvency condition. These studies include Raybaudi et. al.(2004), Matsubayashi (2004), Chortareas et al (2004), Baharumshah et al (2003), Arize, (2002), Wu and et al (2001), Irandoust and Sjöö (2000, 2004), Apergis et al. (2000), Leachman and Francis (2000), Fountas and Wu (1999), Yan (1999), Bodman (1997), Wu and et al (1996), Liu and Tanner (1996, 1995), Sawada (1994), Husted (1992), Hakkio and Rush (1991), among others. As a result of above researches sustainability of external debt, a stationary current account is consistent with a finite external debt-to-GNP ratio. Thus, there is no condition to encourage the country to default on its international debts (Wu, 2000).

Furthermore, the stationarity of the current account is also essential for the validity of the modern intertemporal model of the current account. Theoretically, to interpret the current account acting as prevention to smooth consumption in the face of shocks, the modern intertemporal model of current account determination combines the assumptions of perfect capital mobility and consumption-smoothing behaviour. This emphasizes that the current account series should be a stationary series. Empirically, literature states the Campbell and Shiller (1987) technique to test the validity of the intertemporal model of the current account (e.g. Otto (1992), Ghosh (1995), Shibata and Shintani (1998)). Stationarity tests on current accounts were utilized for the US by Trehan and Walsh (1991), Wickens and Uctum (1993) and for the US and Canada by Otto (1992), and for twenty-three countries by Gundlach and Sinn (1992). These articles state the evidence that current accounts are nonstationary for many industrialized countries, including the US, the UK, Canada, Germany, and Japan. But a common feature of these articles is the results found: a nonstationary current account using the ADF unit-root tests and the other unit-root tests. Wu (2000) using the panel data unit-root test of Im et al. (1997) has re-examined the time-series property of current account among industrial countries. His empirical findings support stationarity of current account series of these countries. This also leads to support the modern intertemporal model of the current account approach.

Additionally, these findings also state that the current account deficits among industrialized countries are sustainable. Sustainability tests, however, do not provide a consensus because results vary with the approach adopted, the sample period, the specification of the transversality condition, and the econometric methodology used (Chortareas et al. 2004). For example, Gundlach and Sinn (1992) analyse time series data and reject a non-stationary current account for most of OECD countries excepting Germany, Japan and the US. Taylor (1996, 2002) analyses historical data series for a sample of 12 countries, which include Germany, Japan and US, and establishes current account stationarty for the entire sample. Chortareas et al. (2004) used a new methodology (Non-Linear Stationarity Tests) that supports sustainability in the debt of a set of Latin America countries.

Conventional unit-root tests have low power when the root is close to one. In parallel, Shiller and Perron (1985) find that the power of the ADF unit-root tests is low with short time spans. Hence, a main reason for the failure of rejecting the non-stationarity of the current account might be the low power of the test in the existing literature. Moreover, Diebold and Rudebusch (1991) and Sowell (1990) found out that conventional unit-root tests such as the Dickey-Fuller test could have low power against fractional alternatives. The traditional views of modelling current account series either as trend deterministic $I(0)$ or as stochastic trends (or unit roots), $I(1)$ processes seem too restrictive compared to the wide scope of possibilities covered by the fractionally integrated $I(d)$ processes. These processes belong to a broader class called long memory, owing to their ability to display significant dependence between distant observations in time (Gil-Alana, 2002a). Thus, the time series properties of current account of G-7 countries are analyzed in this article using the fractional methods, such as Lo's R/S method, Robinson's score test and Whittle's approximate maximum likelihood estimator.

The purpose of this study is to examine the behaviour of current account of seven developed countries for 1974:1-2001:3 time period using fractional methods. The study concentrates on the following major points: First, we determine whether current account series reflect a long memory process. If a current account series is covariance nonstationary but long memory process, then it is classified as mean-reverting process. Thus, the shock has no permanent effect on the values of series. Determining whether the current account series are mean-reverting process is a prerequisite for the analysis of these series. Second, if we view the current account as the series realization of a stochastic process, the autocorrelation function exhibits persistence which is neither consistent with unit root process nor stationary process. Fractionally integrated process is related to the rate of decay associated with the impulse response coefficients of a process. Thus, we estimate the fractionally integrated autoregressive moving average model for each current series for G-7 countries and analyze the calculated impulse response coefficients based on these estimated models. The major findings of this study are: First, all countries' current accounts are covariance non-stationary. Second, three-countries' (France, Italy and Canada) current accounts are mean reverting so that they are sustainable, four-countries' (Germany, UK, US, and Japan) current accounts are not mean reverting, and therefore unsustainable.

The paper is organized as follows: Section II contains some economic foundations of current account sustainability. Section III provides econometric methodology utilized in analysis. Empirical results are reported and interpreted in Section IV. Finally, Section V contains some concluding remarks.

2. A Framework for Testing

The recent literature on the subject under investigation has established the time-series implications of the intertemporal solvency condition. The key result of this literature is that if a country's intertemporal budget constraint is satisfied, then the country is solvent and its path of current account imbalances becomes sustainable. The standard theoretical criterion for assessing current account imbalance is the notion of solvency: a country is solvent to the extent that the discounted value of the expected stock of its foreign debt in the infinitely distant future is non-positive. One of the most common univariate approaches is taken by Liu and Tanner (1996). This approach starts from the per-period budget constraint faced by the country expressed in real terms:

$$M_t - X_t + r_t B_{t-1}^f = \Delta B_{t-1}^f \qquad (1)$$

where M_t is imports in period t, X_t is exports in period *t*, B_t^f is the stock of one period foreign debt issued in period *t* and r_t is the one-period world interest rate. Assuming that the interest rate is stationary with mean *r (r1 = r + vt ,, with vt a zero-mean random error)*, forward iteration of (1) gives

$$B_t^f = \sum_{i=1}^{\infty} \frac{M_{t+1+i} - X_{t+1+i}}{(1+r)^{i+1}} + \lim_{i \to \infty} \frac{B_{t+1+i}^f}{(1+r)^{i+1}} + \sum_{i=1}^{\infty} \frac{v_{t+1+i}}{(1+r)^{t+1}} \tag{2}$$

Now, assuming that exports and imports series are I(1) and taking expected values, equation (2) may be written as;

$$CA_t = \theta + \lim_{i \to \infty} E_t \left[\frac{B_{t+k+i}^f}{(1+r)^{k+1}} \right] + \omega_t, \tag{3}$$

where ω_t is a stationary error term and θ is a constant. This is the no Ponzi condition, in which external debt repayments are sustainable, or the current debt must be equal to the expected present value of future current account surpluses. If the term in the limit was negative, the economy would be consuming and investing more than present value of its future current account surpluses that never converges to zero. If the term in the limit was greater than zero, the country would be paying the old maturity debt by issuing new debt, which reveals that current account is not sustainable in the long-run. If the solvency condition holds, then the second term on the right hand side of equation (3) is equal to zero. Therefore, the solvency condition requires that current account must be a stationary process around a constant mean (see Trehan and Walsh (1991)).

An alternative way of testing the solvency condition requires investigating the cointegration of exports and imports. Indeed, Liu and Tanner (1996, p.741) emphasize three important advantages of the stationary test. "First, the stationarity test is somewhat stronger than the cointegration test, as it imposes a vector of cointegrating coefficient of [1, -1] on the variables exports and imports. Second, as Trehan and Walsh (1991) show, while the interest rate must be stationary for the cointegration test, it need not be so for the stationarity test. Third, in the stationary test, exports and imports need not follow random walks for the test to be valid. For these reasons, the stationarity test is more appropriate than the cointegration test."

The stationarity of the current account is important to the validity of the intertemporal model of the current account. Theoretically, this approach combines the assumptions of perfect capital mobility and consumption smoothing behavior to predict that current account acts as a buffer to smooth consumption in the face of shocks and implies that current account will typically be a stationary variable. Therefore, the solvency constraint requires that the current account be a stationary variable. This means that current account is mean-revert. In this paper, we also study the order of integration of current account using a methodology based on fractional integration. In the unit-root tests, the knife-edge distinction between $I(0)$ and $I(1)$ processes of the current account can be too restrictive. If the process is $I(1)$, it is not mean-revert and the effect of shock is persistent. In the long-memory models, if the degree of integration of current account is within (0, 0.5) interval, current account is covariance stationary and is mean revert. If the degree of integration of current account is within (0.5,1) interval, the current account is not covariance stationary, but it is mean reverting, with the effect of shocks dying away in the long run. Finally, if the $d \geq 1$, current account is nonstationary and non-mean reverting. In this context, the current account deficit sustainability condition holds if and only if the fractional integration parameter of the current account series is less than unity. There exist many approaches of estimating and testing the fractional differencing parameter d (e.g. Geweke and Porter-Hudak, 1983).

3. Econometric Methodology

Over the last two decades, there has been a growing literature modelling the macroeconomic and financial time series in terms of fractionally integrated processes. These processes constitute a generally class of flexible time series models. They are useful for modelling low-frequency dynamics, because of their flexibility and simplicity in their specification (Granger and Joyeux, 1980; Hosking, 1981).

Let y_t, $t = 1, 2, \ldots, T$, be the time series of interest. In the frequency domain, assume that yt is weakly stationary process and its spectral density function, $f(\lambda)$, at frequency $\lambda \in (-\pi, \pi]$ satisfying

$$\gamma_j = E[(y_t - E(y_t))(y_{t+j} - E(y_t))] = \int_{-\pi}^{\pi} f(\lambda) \cos(j\lambda) d\lambda, \quad j = 0, \pm 1, \pm 2, \ldots \qquad (4)$$

where λ_j are the autocovariance of y_t. Spectral density function of y_t satisfy

$$f(\lambda) \sim c_1 \lambda^{-2d} \qquad \text{as } \lambda \to 0^+ \qquad \text{for } 0 < c_1 < \infty \qquad (5)$$

and autocovariances follow

$$\gamma_j \sim c_2 j^{2d-1} \qquad \text{as } j \to \infty \qquad \text{for } |c_2| < \infty \qquad (6)$$

where the symbol ~ means that the ratio of the left hand side and right hand side tends to 1, as $j \to \infty$ in equation (6) , and as $\lambda \to 0^+$ in equation (5). For $d \in (-0.5,0.5)$, y_t follows a long memory process. Condition equation (5) and equation (6) are not always equivalent, but Yong (1974) and Zygmund (1995) give conditions under which both expressions are equivalent (Brockwell and Davis, 1991; Robinson, 1995a,b; Baillie, 1996).

A general class of fractional integrated processes ARFIMA (*p*,*d*,*q*) is described by

$$\phi(L)\,(1-L)^d\, y_t = \theta(L)\varepsilon_t \tag{7}$$

where L is lag operator, $\phi(L) = 1 - \phi_1 L - \ldots - \phi_p L^p$ and $\theta(L) = 1 - \theta_1 L - \ldots - \theta_q L^q$ are polynomials with stable roots, ε_t is white noise, d is long memory parameter and the fractional differencing operator, $(1-L)^d$, yields an infinite-order lag polynomial with slowly declining coefficients as bellows:

$$(1-L)^d = \sum_{k=0}^{\infty} \Gamma(k-d)L^k \,/ \{\Gamma(k+1)\Gamma(-d)\}$$

where $\Gamma(\cdot)$ is the standard gamma function. These processes belong to long memory due to their ability to reveal significant dependence between distant observations in time. The order of integration d is not restricted to integer values and can take any value on the real line. Unit root processes are included as a special case with $d = 1$. Stationary case is with $d = 0$. For $d \in (0,0.5)$, y_t is said to have long memory. When $d \in (-0.5,0)$, y_t is called antipersistent or intermediate memory. For $d \le -0.5$, y_t is covariance stationary but not invertible. For $d \ge -0.5$, y_t is nonstationary and has infinite variance. For $d \in (0.5,1)$, where y_t displays strong persistence, but mean reverts in the sense that the impulse response function is decaying (Granger and Joyeux, 1980; Hosking, 1981).

The mean-reverting property depends on whether $d < 1$. The impact of a shock is known to be persistent forever when $d = 1$. The effect of any shock on the fractionally integrated process with $d < 1$ slowly dies out. This can be seen by studying the moving average representation for $(1-L)y_t$ as follows:

$$(1-L)y_t = A(L)\varepsilon_t \tag{8}$$

where $A(L) = (1+\theta_1 L+\theta_2 L^2+\ldots)$ obtained from $(1-L)^{1-d}\Phi(L)$ with $\Phi(L) = \phi^{-1}(L)\theta(L)$. The impulse responses are given by the coefficients θ_k of $A(L)$. The impact of a unit innovation at time t on the value of y_t at $t+j$ is equals to equation 9.

$$C_j = (1+\theta_1+\theta_2+\ldots+\theta_j) \tag{9}$$

In equation 9, as $j \to \infty$, $C_\infty = A(1)$. That is the measure of the long run impact of the innovation (Campbell and Mankiw, 1987). Cheung and Lai (1993) show that for the fractionally integrated process with $d < 1$, $C_\infty = 0$ implying no long-run impact of the innovation on the value of y_t. For $d \geq 1$, $C_\infty \neq 0$. So, the effect of a shock has permanent effect on the value of y_t. When the $d < 1$, the y_t process is mean-reverting (Cheung and Lai, 1993; Cheung and Lai, 2001).

Testing Procedures for Fractional Integration

Hurst (1951) developed "rescaled range" statistic (R/S) detecting evidence of strong dependence in time series. It is popularised by Manderbold (1972). Lo (1991) shows that short-range dependence may compromise inferences about the presence of long-range dependence. He derives an adjustment to the classical R/S statistics accounting for general forms of short-range dependence. The modified R/S statistics replaces the usual variance estimate with a consistent estimator of the "long run variance". Cavaliere (2001) proposes a new generalized R/S statistics. Robinson (1994a) proposes a very general testing procedure for testing unit root and other hypothesis in raw time series. Unlike most of unit root tests embedded in autoregressive alternatives, Robinson's score test is nested in a fractionally integrated model. Le and Schmidt (1996) propose the test of Kwiatowski, Phillips, Schmidt and Shin (1992), KPSS, as a test for the null of stationarity against the alternative hypothesis of fractional integration. Dolado, Gonzalo and

Mayoral (2002) proposed a Fractional Dickey-Fuller test (FDF) for testing the null hypothesis of $I(d_0)$ against the alternative hypothesis $I(d_1)$, $d_1 < d_0$. In this paper, we use the Lo's modified R/S statistics and Robinson's score test statistics detecting or testing the long memory in current account series of G-7 countries. Also, reduced form of Whittle estimator is used for the long memory models. Thus, these are briefly outlined in this section.

Lo's Modified R/S Statistics

The modified R/S test for long memory considers the null hypothesis of a short memory against the alternative hypothesis of long memory. Let $\bar{y}$ be the sample mean of time series $\{y_1, y_2, \ldots y_T\}$. The modified R/S statistics, $Q_{n,q}$, is given by the range of cumulative sums of derivations of time series from its mean, rescaled by a consistent estimate of its standard deviation:

$$Q_{n,q} = S_q^{-1}\{\max_{1\le k\le T}\sum\nolimits_{j=1}^{k}(y_j - \bar{y}_n) - \min_{1\le k\le T}\sum\nolimits_{j=1}^{k}(y_j - \bar{y}_n)\} \tag{10}$$

where S_q^2 is a heteroskedasticity and autocorrelation consistent variance estimator (Andrews, 1991),

$$S_q^2 = \left\{\sum\nolimits_{i=1}^{T}(y_i - \bar{y})^2 / T + 2\sum\nolimits_{j=1}^{q}\tau_j(q)\left(\sum\nolimits_{i=j+1}^{T}(y_i - \bar{y})(y_{i-j} - \bar{y})\right)/T\right\}$$

with the weighting function $\tau_j(q) = 1 - |\, j / z_T \,|$ and a truncation lag q determined by

$$q = \mathrm{Int}[z_T]_{,}\ z_T = (3T/2)^{1/3}\{2\rho/(1-\rho^2)\}^{2/3},$$

where $\mathrm{Int}[z_T]$ denotes the integer part of the z_T and ρ is the first-order autocorrelation of the series.

The modified R/S statistics is different from the classical one on the normalization of the range measure. The denominator in equation (10) normalizes the range measure both by the sample variance (q = 0) (as considered in classical R/S analysis) and by a weighted sum of sample autocovariances for $q > 0$. Such modification contributes to the robustness of the modified R/S analysis to take care of short-term dependence and heteroskedasticity. The limiting distribution of the $Q_{n,q}$ statistic standardized by the square root of the sample size may be established under the null hypothesis of no long memory. Lo (1991) provides the critical values for the modified R/S tests.

Robinson LM Test

Robinson (1994a) proposes a very general procedure for testing unit roots as well as other nonstationary alternatives. Unlike the other unit root tests (e.g. Dickey and Fuller, 1979; etc.), testing for autoregressive (AR) unit root, Robinson's (1994a) procedure allows testing for fractional order of integration in addition to other appealing hypothesis. Robinson's (1994a) tests are nested in a fractionally integrated model. Denoting the y_t time series, we employ throughout the model,

$$y_t = \beta' z_t + x_t, \quad t = 1, 2, \ldots, \tag{11}$$

$$(1-L)^d x_t = u_t, \quad t = 1, 2, \ldots, \tag{12}$$

where y_t is time series we observe for $t = 1, 2, \ldots, T$, β is a (k×1) vector of unknown parameters; z_t is a (k×1) vector of deterministic regressors, such as polynomials in *t* (usually an intercept and a linear time trend), x_t is the regression error, *d* is a given real value and u_t is a covariance stationary process with spectral density function which is positive and finite at the zero frequency.

Robinson (1994a) proposes a score test of the null hypothesis, defined by:

$$H_0 : d = d_0. \tag{13}$$

The score statistic for above testing the hypothesis proposed by Robinson (1994a) is given by

$$\hat{r} = (T / \hat{A})^{0.5} (\hat{a} / \hat{\sigma}^2), \tag{14}$$

where T is the sample size and

$$\hat{a} = (-2\pi / T)\sum_{j=1}^{T-1}\psi(\lambda_j)g(\lambda_j;\hat{\tau})^{-1}I(\lambda_j);\ \ \hat{\sigma}^2 = (2\pi / T)\sum_{j=1}^{T-1}g(\lambda_j;\hat{\tau})^{-1}I(\lambda_j);$$

$$\hat{A} = (2/T)\left(\sum_{j=1}^{T-1}\psi(\lambda_j)^2 - \sum_{j=1}^{T-1}\psi(\lambda_j)\hat{\varepsilon}(\lambda_j)' \times \left(\sum \hat{\varepsilon}(\lambda_j)\hat{\varepsilon}(\lambda_j)'\right)^{-1} \times \sum_{j=1}^{T-1}\hat{\varepsilon}(\lambda_j)\psi(\lambda_j)\right)$$

$$I(\lambda_j) = \frac{1}{2\pi T}\left|\sum_{t=1}^{T}\hat{u}_t e^{i\lambda_j t}\right|^2; \psi(\lambda_j) = \log\left|2\sin(\lambda_j / 2)\right|; \hat{\varepsilon}(\lambda_j) = \frac{\partial}{\partial\tau}\log g(\lambda_j;\hat{\tau}); \lambda_j = (2\pi j / T).$$

where $I(\lambda_j)$ is the periodogram of $\hat{u}_t$ obtained as:

$$\hat{x}_t = y_t - \hat{\beta}z_t,\ \hat{\beta} = \left(\sum_{t=1}^{T} z_t z_t'\right)^{-1}\sum_{t=1}^{T} z_t y_t,\ \hat{u}_t = (1-L)^d \hat{x}_t,\ t = 1,2,\ldots,$$

and g above is a given function coming from the spectral density function of $\hat{u}_t$: $f(\lambda;\tau) = (\sigma^2 / 2\pi)g(\lambda;\tau)$, with $\hat{\tau}$ obtained by minimising $\sigma^2(\tau)$. If u_t is white noise, $g \equiv 1$. If u_t is an AR process from: $\phi(L)u_t = \varepsilon_t$, $g = \left|\phi_p(e^{i\lambda})\right|^{-2}$. Therefore, AR coefficients are function of τ.

Under the null hypothesis given in (13) Robinson (1994a) established that

$$\hat{r} \rightarrow_d N(0,1) \quad \text{as} \quad T \rightarrow \infty, \tag{15}$$

under the certain regularity conditions. This limiting distribution holds independently of the regressors included in z_t and the various types of I(0) disturbances assumed for u_t including the general weakly stationary in (12). An approximate one-sided test of $H_0 : d = d_0$ is rejected against the alternative: $H_a : d > d_0$ $(d < d_0)$ at the α % level will be given by the rule "Reject H_0 if $\hat{r} > z_\alpha$ $(\hat{r} < -z_\alpha)$", where the probability of a standard normal variate exceeding z_α is α. The above version of the Robinson test is used in empirical applications by in Gil-Alana and Robinson (1997) and Gil-Alana (1999, 2000, 2001, 2002b).

Estimation method for long memory models

In this paper, we will evaluate the persistence for testing of intertemporal solvency using the impulse response functions of the estimated ARFIMA models. In order to obtain impulse responses, we first need to estimate parameters of the models. There exist different approaches for estimating parametric models like equation (7) and analysis may be carried out in the frequency and in the time domain. Fox and Taqqu (1986) proposed a frequency domain method to estimate ARFIMA models by minimizing the Whittle function (that is an approximation to the exact likelihood function). Dalhaus (1989) proposed another method to estimate ARFIMA models by minimizing the exact likelihood function in frequency domain. Sowell (1992) analyzed in the time domain the exact maximum likelihood (EML) estimates of the parameters of an ARFIMA model using recursive procedures allowing quick evaluation of the likelihood function. As Hauser (1999) motioned these estimators are asymptotically equivalent, but Whittle maximum likelihood (WML) with respect to the EML estimator is more reliable in small samples. Thus, WML estimator is used in this paper. The WML estimates are obtained by maximizing an approximation of the likelihood function of the ARFIMA model in equation (7) in the frequency domain. In this method, the parameter vector $\theta = (\alpha 1,\dots,\alpha p,d,\beta 1,\dots,\beta q)$ is estimated by minimizing the following approximate log likelihood function

$$\log L_W(\theta,\sigma_u^2) = -\sum_{j=1}^{m} \log f(\lambda_j \mid \theta,\sigma_u^2) - \frac{1}{2\pi}\sum_{j=1}^{m}\frac{I(\lambda_j)}{f(\lambda_j \mid \theta,\sigma_u^2)} \tag{16}$$

where $I(\lambda j)$ is the periodogram defined at the j^{th} Frourier frequency, $\lambda_j = 2\pi j/T, j = 1,\dots,m,$

$$I(\lambda_j) = T^{-1}\left|\sum_{t=1}^{T}(y_t - \bar{y})e^{it\lambda_j}\right|^2 \tag{17}$$

$m = [(T-1)/2]$, $[\cdot]$ is the integer part. The reduced form of L_W with respect to the error variance σ_u^2 is

$$\log L_W^*(\theta) = m\log(2\pi) - m\log\left[\frac{1}{m}\sum_{j=1}^{m}\frac{I(\lambda_j)}{g(\lambda_j)}\right] - \sum \log g(\lambda_j) - m \qquad (18)$$

with $\sigma_u^2 = \sigma_u^{2*} = m^{-1}\sum_{j=1}^{m}(I(\lambda_j)/g(\lambda_j))$ where $f(\lambda_j) = \sigma_u^2 g(\lambda)/(2\pi)$ with $g(\lambda) = g(\lambda \mid \theta)$. In this paper, the parameters of each ARFIMA(*p,d,q*) model for real exchange rate series are estimated by reduced form of L_W (Hauser, 1999).

IV. Data and Empirical Results

The quarterly data are taken from International Monetary Fund's Balance of Payments Statistics (CD-ROM, Version Apr. 2002) for the 1974:1- 2001:3 period. The series are seasonally adjusted using X-11 procedure.

The modified R/S test is performed on current account series, and the results are reported in Table 1. The modified R/S test statistics are provided together with the optimal lag selected and used in constructing the corresponding statistics. The results for the current account series for G-7 countries show that in all cases the null hypothesis of short-memory can be rejected at the 5 percent significance level. This finding shows the presence of long memory in current account for G-7 countries.

Table 1: Results of modified R/S analysis for current account series

Country	R/S Statistic	q-Lag selected
Canada	3.081*	0
Germany	1.955*	3
Italy	2.109*	1
USA	2.555*	1
UK	2.205*	2
France	1.909*	3
Japan	1.878*	2

Notes: The lag parameter q used for the modified R/S test is determined by Andrews's (1991) data-dependent rule. At the 5% significant level, the null hypothesis of a short-memory process is rejected if the modified R/S statistics does not fall within the confidence interval [0.809, 1.862]. * shows that the null hypothesis is rejected at the 5 percent significant level.

In Appendix A, table 1, 2, 3, 4, 5, 6, and 7, report the values of the one-sided test statistics $\hat{r}$ of equation (14). The significant negative values of this statistics are consistent with $H_a : d < d_0$, whereas the significant positive values are consistent with $H_a : d > d_0$. Thus, a monotonic decrease in the value of the $\hat{r}$ statistic is expected with increasing *d* values, because if the null hypothesis is rejected for a certain value of *d*, then a more significant result should be expected for testing values than this value of *d*. In this study, we consider two cases for the Data Generating Process (DGP) for u_t. In the first case, we assume that u_t is white noise. In the second case, u_t is assumed to be follow an AR(1) process. In these tables, we present the results based on various z_t including $z_t = 0$ (i.e., including no regressors in the undifferenced regression), $z_t = 1$ (i.e., including an intercept) and $z_t = (1,t)'$ (i.e., including an intercept and a linear time trend). In all cases, the values of the $\hat{r}$ statistics are monotonically decreasing with increasing *d* values. Thus, the model is not likely to be misspecified. The first rows of tables give the values of *d* considered under the null hypothesis in equation (13). The next rows present the results for $z_t = 0$, $z_t = 1$ and $z_t = (1,t)'$. The first panel of Table 1 reports the values for white noise disturbances, while in second panel of Table 1 we allow AR(1) disturbances for Canada. In the first panel of Table 1, if the current account series of Canada is modelled with $z_t = 0$, $z_t = 1$ and $z_t = (1,t)'$, the range of d values is accepted by the Robinson test at the five percent significant level between 1.05-1.30, 0.95-1.25 and 0.90-1.15, respectively. If an AR(1) process is assumed for the disturbances u_t with $z_t = 0$, $z_t = 1$ and $z_t = (1,t)'$, the range of d values accepted by the Robinson test is between 1.10-1.20, 1.05-1.15 and 1.00-1.10, respectively.

It is observed that, in the Table 2, starting with the case of white noise disturbance, null hypothesis cannot be rejected when the range of *d* values is between 0.95 and 1.15. However, if we allow weakly parametrically autocorrelated disturbances (AR(1) disturbance case) along with the $z_t = 0$, $z_t = 1$ and $z_t = (1,t)'$ components, then the order of integration values accepted by Robinson test is between 1.00 and 1.05 for $z_t = 0$ and $z_t = 1$, respectively, and is 1.00 for $z_t = (1,t)'$. If we look at Table 3, results of Italy show us that the non-rejection values of d range from 0.90 to 1.10 for $z_t = 0$ and $z_t = (1,t)'$, respectively, and range from 0.90 to

1.15 for $z_t = 1$ in case of white noise disturbances; between 0.95 and 1.00 with AR(1) disturbances.

Table 4 shows that if u_t is white noise with $z_t = 0$ and $z_t = 1$, non-rejection values range between 1.15 and 1.30 for USA. On the other hand, for the $z_t = (1, t)'$, the order of integration is between 1.20 and 1.30. When the disturbances are assumed to follow an AR(1), the order of integration is between 1.20 and 1.30 for three different deterministic regressors. Results of U.S. show that current account series of U.S. is not sustainable.

Table 5 presents the results for U.K. For the result given in the first panel in the Table 6 the range of values for that the null hypothesis for $z_t = 0$, $z_t = 1$ and $z_t = (1, t)'$ cannot be rejected is between 0.80 and 1.00. The second panel shows that the range of values d that is not rejected at the five percent significance level under the null hypothesis is between 0.85 and 0.90. Except d = 1.00 case for U.K., the current account series of U.K. can be represented by a covariance nonstationary and long memory process. Thus, the current account series is mean-reverting series showing that it is sustainable. But, it is not sustainable in the latter case.

The results given in first panel of the Table 6 for France are obtained by assuming a white noise process for the disturbances with $z_t = 0$, $z_t = 1$ and $z_t = (1, t)'$ deterministic regressors. For the first two deterministic cases (no deterministic regressors and an intercept case) and last deterministic case, the range of values for which the null hypothesis cannot be rejected is from 0.80 to 1.00 and from 0.75 to 1.00, respectively. In the case where we assume an AR(1) process for disturbances and the same set of deterministic regressors, we see that the range of d values accepted by the Robinson test is between 0.85 and 0.90.

For Japan, if we consider the white noise u_t and these deterministic regressors, we can see that degrees of integration of current account series is bigger than 1.30. If we allow an AR(1) process for u_t along with these deterministic regressors, we see that the nonrejection value of d is 1.30 in case of no deterministic regressors; ranges between 1.20 and 1.30 with an intercept; ranges between 1.25 and 1.30 with an intercept and a linear time trend. These results show that current account series of Japan is not sustainable.

For each series we present the estimating results different ARFIMA(*p,d,q*) models where both *p* and *q* are less than or equal to three. The selection of the best ARFIMA model is based on the Schwarz information criterion (SIC). For the p,q > 3, different ARFIMA models are also estimated. But, estimation of these models either failed or, when estimation was successful, they always had lower SIC values. Therefore, we do not consider these higher order ARFIMA models. For brevity only the results of ARFIMA model estimates having the minimum SIC will be presented. The parameters of the ARFIMA models are estimated using the WML method. Standard errors for the ARFIMA estimates are calculated using the asymptotic formula in Robinson (1994b) and Beran (1995).

The estimation results for the parametric ARFIMA models for each country are given in Table (2). Based on the minimum SIC criteria, the best ARFIMA model for each country is an ARFIMA(3,0.67,3) for the Canada, an ARFIMA(1,1.09,0) for Germany, an ARFIMA(2,0.57,0) for Italy, an ARFIMA(0,1.05,0) for USA, an ARFIMA(2,1.06,0) for UK, an ARFIMA(2,0.54,3) for France and an ARFIMA(0,1.14,0) for Japan, respectively. The estimated fractional differencing parameter of each current account series of each country exhibits fractional dynamics with long-memory features. They range from 0.54 to 1.14 for the series under consideration. The estimated values of d are significantly different from zero at the 5% level for each country. Therefore, each series is not covariance stationary and exhibits long-memory behaviour. When compared to the estimated value of d of these series, the estimated value of *d* for Germany, USA, UK and Japan is larger than 1. Hence, lower band of confidence intervals for long memory parameters of these series should be checked. Calculated confidence intervals show that lower bands of these are 0.91, 0.91, 0.84 and 0.98, respectively. Also, this indicates that these current accounts are not explosive series.

Table 2: Parameter Estimates of Best ARFIMA(p,d,q) models for CA

Country	log-lik	d	α1	α2	α3	β1	β2	β3	SIC
Canada	57.18	0.67 (0.11)	1.10 (0.73)	-0.59 (0.99)	-0.23 (0.69)	0.91 (0.64)	-0.47 (0.84)	-0.39 (0.64)	-98.778
Germany	48.98	1.09 (0.09)	-0.33 (0.12)	-	-	-	-	-	-91.005
Italy	47.53	0.51 (0.31)	0.32 (0.31)	0.31 (0.10)	-	-	-	-	-86.409
USA	45.04	1.05 (0.07)	-	-	-	-	-	-	-84.885
UK	52.75	1.06 (0.11)	-0.37 (0.13)	-0.29 (0.11)	-	-	-	-	-96.855
France	53.61	0.54 (0.15)	0.68 (0.25)	-0.46 (0.22)	-	0.61 (0.27)	-0.58 (0.24)	-0.21 (0.16)	-93.508
Japan	40.99	1.14 (0.08)	-	-	-	-	-	-	-76.941

The estimates given in the Table are for the models that have the minimum SIC. The values in parentheses are standard errors of parameters. Standard errors are calculated under the asymptotic formula in Robinson (1994b) and Beran (1995).

As an alternative confirmation of the solvency of current account, we analyse the impulse responses implied by the ARFIMA models selected by the SIC for each country. Persistence of current account dynamics can be analysed through the sequence of C_j, as given by equation (9). The function, that gives the sequence of C_j values at different time horizon after a unit shock, can be computed based on equation (8) for current account series. In Appendix A, Figure (1) displays the plots of the impulse responses for the selected model in each country when they are shocked by one standard deviation. Graphs of the first 20 dynamic responses show different dynamics. For France, Italy and Canada, adjustment process has zero long-run persistence confirming that current account series of these countries are sustainable. However, for the Canadian case, the graph exhibits nonlinearity. In this case, the process of adjustment shows nonlinearity in the direction of adjustment because of cyclical responses to the initial shock. In Germany, UK, USA and Japan, adjustment process has not zero long-run persistence. In other words, current account series of these countries are not mean-reverting indicating that these are unsustainable.

V. Concluding Remarks

In this article we have examined the sustainability of G-7 current account by means of fractional integration techniques. Using a version of the Robinson (1994) for testing unit and fractional roots, the results show that all countries current account is covariance non-stationary and four-country (France, UK, Italy and Canada) current accounts are mean reverting so that they are sustainable. Germany, US, and Japan's current account are not mean reverting and hence are unsustainable.

The results here reinforce the existence of mean reversion in the current account, though, in view of the values of *d*, ranging in most cases between 0.5 and 1, the adjustment process towards equilibrium will take a very long time. The persistence graphs show different dynamics. For France, Italy and Canada, adjustment process has zero long-run persistence confirming that current account series of these countries are sustainable. However, for the Canadian case, the graph exhibits nonlinearity. In Germany, UK, USA, and Japan adjustment process has not zero long-run persistence. In other words, current account series of these countries are not mean-reverting indicating that these are unsustainable.

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Appendix A

Table 1: CANADA

Testing the order of integration with the tests of Robinson and white noise disturbances

Z_t/d	0.00	0.05	0.10	0.15	0.20	0.25	0.30	0.35	0.40	0.45	0.50	0.55	0.60	0.65	0.70	0.75	0.80	0.85	0.90	0.95	1.00	1.05	1.10	1.15	1.20	1.25	1.30
$Z_t = 0$	25.28	24.59	23.72	22.65	21.39	19.97	18.43	16.82	15.21	13.63	12.12	10.69	9.35	8.11	6.94	5.87	4.87	3.96	3.12	2.36	1.67	1.05	0.49	-0.01	-0.46	-0.87	-1.24
$Z_t = 1$	17.18	16.66	16.14	15.61	15.04	14.42	13.74	12.99	12.15	11.24	10.26	9.23	8.16	7.09	6.03	5	4.03	3.13	2.31	1.57	0.91	0.32	-0.2	-0.67	-1.08	-1.44	-1.77
$Z_t = (1,t)^t$	16.45	15.87	15.24	14.54	13.79	12.98	12.11	11.19	10.22	9.22	8.2	7.17	6.17	5.19	4.26	3.38	2.57	1.83	1.15	0.54	-0.01	-0.49	-0.93	-1.31	-1.66	-1.97	-2.25
Testing the order of integration with the tests of Robinson and AR(1) disturbances																											
$Z_t = 0$	372.67	322.99	273.09	226.36	184.48	148.22	117.8	92.96	73.15	57.6	45.5	36.13	28.85	23.18	18.7	15.14	12.24	9.82	7.76	5.95	4.31	2.81	1.42	0.14	-1.04	-2.12	-3.1
$Z_t = 1$	184.11	162.75	141.47	121.41	103.24	87.2	73.22	61.16	50.83	42.07	34.69	28.52	23.39	19.14	15.59	12.61	10.08	7.89	5.97	4.25	2.69	1.27	-0.02	-1.19	-2.25	-3.21	-4.08
$Z_t = (1,t)^t$	136.42	120.35	104.64	89.9	76.48	64.55	54.12	45.14	37.49	31.01	25.58	21.02	17.22	14.02	11.33	9.02	7.02	5.26	3.67	2.23	0.91	-0.31	-1.43	-2.46	-3.4	-4.25	-5.01

Notes: * and in bold indicate non-rejection values at the five percent significance level.

Table 2: GERMANY

Testing the order of integration with the tests of Robinson and white noise disturbances

Z_t/d	0.00	0.05	0.10	0.15	0.20	0.25	0.30	0.35	0.40	0.45	0.50	0.55	0.60	0.65	0.70	0.75	0.80	0.85	0.90	0.95	1.00	1.05	1.10	1.15	1.20	1.25	1.30
$Z_t = 0$	22.13	21.35	20.5	19.6	18.64	17.63	16.57	15.46	14.3	13.11	11.87	10.61	9.33	8.06	6.82	5.63	4.49	3.43	2.45	1.56*	0.76*	0.04*	-0.6*	-1.16*	-1.65	-2.09	-2.47
$Z_t = 1$	18.83	18.44	18.01	17.51	16.94	16.29	15.54	14.7	13.76	12.73	11.61	10.43	9.21	7.96	6.72	5.52	4.37	3.29	2.3	1.39*	0.58*	-0.14*	-0.78*	-1.35*	-1.84	-2.27	-2.65
$Z_t = (1,t)^t$	19.05	18.67	18.23	17.72	17.12	16.44	15.66	14.78	13.81	12.75	11.6	10.39	9.14	7.87	6.62	5.4	4.23	3.14	2.13	1.22*	0.4*	-0.33*	-0.97*	-1.54*	-2.03	-2.46	-2.83
Testing the order of integration with the tests of Robinson and AR(1) disturbances																											
$Z_t = 0$	268.01	235.33	203	172.34	144.32	119.46	97.9	79.54	64.18	51.52	41.24	33	26.43	21.2	16.98	13.49	10.5	7.82	5.34	3.03	0.91*	-0.98*	-2.64	-4.05	-5.26	-6.29	-7.17
$Z_t = 1$	188.31	173.44	157.18	139.92	122.25	104.86	88.38	73.35	60.13	48.84	39.47	31.81	25.64	20.65	16.57	13.15	10.16	7.46	4.95	2.62	0.48*	-1.42*	-3.07	-4.48	-5.67	-6.7	-7.58
$Z_t = (1,t)^t$	186.79	172.28	155.95	138.43	120.49	102.95	86.5	71.64	58.68	47.69	38.59	31.16	25.15	20.27	16.25	12.85	9.86	7.13	4.59	2.23	0.08*	-1.83	-3.48	-4.9	-6.1	-7.12	-8.01

Notes: * and in bold indicate non-rejection values at the five percent significance level.

Table 3: ITALY

Testing the order of integration with the tests of Robinson and white noise disturbances

Z_t/d	0.00	0.05	0.10	0.15	0.20	0.25	0.30	0.35	0.40	0.45	0.50	0.55	0.60	0.65	0.70	0.75	0.80	0.85	0.90	0.95	1.00	1.05	1.10	1.15	1.20	1.25	1.30
$Z_t = 0$	15.71	15.12	14.5	13.83	13.11	12.34	11.53	10.66	9.76	8.82	7.87	6.91	5.96	5.03	4.13	3.27	2.46	1.7	0.99^{*}	0.34^{*}	-0.25^{*}	-0.79^{*}	-1.27^{*}	-1.71	-2.1	-2.45	-2.76
$Z_t = 1$	15.34	14.88	14.37	13.8	13.17	12.49	11.75	10.96	10.11	9.22	8.3	7.36	6.41	5.46	4.54	3.64	2.79	1.99	1.24^{*}	0.55^{*}	-0.08^{*}	-0.65^{*}	-1.17^{*}	-1.63^{*}	-2.04	-2.41	-2.74
$Z_t = (1,t)^t$	16.65	15.81	14.99	14.17	13.33	12.47	11.59	10.68	9.75	8.8	7.84	6.87	5.92	4.99	4.09	3.24	2.43	1.67	0.97^{*}	0.32^{*}	-0.27^{*}	-0.8^{*}	-1.28^{*}	-1.71	-2.1	-2.45	-2.76
Testing the order of integration with the tests of Robinson and AR(1) disturbances																											
$Z_t = 0$	149.51	132.09	116.28	101.72	88.15	75.44	63.64	52.87	43.32	35.08	28.17	22.5	17.92	14.22	11.21	8.69	6.5	4.5	2.6	0.78^{*}	-0.94^{*}	-2.52	-3.92	-5.15	-6.23	-7.19	-8.06
$Z_t = 1$	130.83	119.73	108.51	97.03	85.36	73.74	62.52	52.08	42.74	34.66	27.88	22.32	17.82	14.18	11.21	8.71	6.53	4.53	2.63	0.81^{*}	-0.91^{*}	-2.48	-3.88	-5.12	-6.2	-7.16	-8.02
$Z_t = (1,t)^t$	110.71	102.3	93.28	83.81	74.13	64.55	55.37	46.86	39.21	32.52	26.79	21.96	17.93	14.56	11.72	9.26	7.06	5.02	3.09	1.23^{*}	-0.51^{*}	-2.11	-3.55	-4.82	-5.93	-6.92	-7.8

Notes: * and in bold indicate non-rejection values at the five percent significance level.

Table 4: USA

Testing the order of integration with the tests of Robinson and white noise disturbances

Z_t/d	0.00	0.05	0.10	0.15	0.20	0.25	0.30	0.35	0.40	0.45	0.50	0.55	0.60	0.65	0.70	0.75	0.80	0.85	0.90	0.95	1.00	1.05	1.10	1.15	1.20	1.25	1.30
$Z_t = 0$	18.15	17.85	17.59	17.34	17.09	16.83	16.53	16.18	15.76	15.27	14.68	14	13.2	12.3	11.29	10.2	9.05	7.85	6.66	5.48	4.36	3.31	2.34	1.47*	0.68	-0.02*	-0.63*
$Z_t = 1$	18.89	18.66	18.39	18.09	17.76	17.37	16.94	16.44	15.87	15.21	14.48	13.65	12.74	11.74	10.67	9.55	8.4	7.24	6.09	4.98	3.92	2.94	2.03	1.21*	0.48*	-0.18*	-0.76*
$Z_t = (1,t)^t$	23.25	23.1	22.92	22.7	22.42	22.09	21.69	21.19	20.6	19.88	19.03	18.03	16.88	15.58	14.15	12.61	11.02	9.41	7.84	6.35	4.96	3.71	2.59	1.6	0.74*	0.00*	-0.65*
Testing the order of integration with the tests of Robinson and AR(1) disturbances																											
$Z_t = 0$	373.64	368.72	361.18	348.75	328.4	298.24	259.93	218.45	179	144.6	116.03	92.82	74.15	59.19	47.23	37.68	30.06	23.95	18.99	14.91	11.46	8.49	5.88	3.56	1.52*	-0.28*	-1.83
$Z_t = 1$	205.79	198.45	192.01	185.56	178.23	169.17	157.85	144.18	128.7	112.33	96.07	80.68	66.68	54.33	43.76	34.94	27.76	21.99	17.38	13.64	10.52	7.8	5.34	3.08	1.03*	-0.78*	-2.36
$Z_t = (1,t)^t$	197.39	188.55	178.17	166.27	153.04	138.87	124.27	109.72	95.67	82.43	70.23	59.24	49.52	41.11	33.93	27.9	22.85	18.64	15.1	12.06	9.4	7.02	4.86	2.89	1.09*	-0.53*	-1.96

Notes: * and in bold indicate non-rejection values at the five percent significance level.

Table 5: UK

Z_t/d	0.00	0.05	0.10	0.15	0.20	0.25	0.30	0.35	0.40	0.45	0.50	0.55	0.60	0.65	0.70	0.75	0.80	0.85	0.90	0.95	1.00	1.05	1.10	1.15	1.20	1.25	1.30
Testing the order of integration with the tests of Robinson and white noise disturbances																											
$Z_t = 0$	22.02	21.27	20.38	19.32	18.12	16.78	15.32	13.77	12.18	10.57	8.99	7.47	6.04	4.72	3.52	2.45	1.5*	0.67*	-0.05*	-0.68*	-1.22*	-1.68	-2.08	-2.42	-2.72	-2.98	-3.21
$Z_t = 1$	18.35	17.74	17.04	16.24	15.33	14.3	13.18	11.97	10.69	9.37	8.03	6.72	5.46	4.27	3.16	2.16	1.26*	0.46*	-0.23*	-0.84*	-1.37*	-1.82	-2.21	-2.55	-2.84	-3.09	-3.31
$Z_t = (1,t)^t$	17.96	17.34	16.63	15.82	14.91	13.9	12.8	11.63	10.39	9.12	7.84	6.58	5.36	4.22	3.15	2.18	1.31*	0.53*	-0.15*	-0.75*	-1.27*	-1.72	-2.11	-2.45	-2.75	-3	-3.23
Testing the order of integration with the tests of Robinson and AR(1) disturbances																											
$Z_t = 0$	235.19	199.87	165.81	134.82	107.91	85.36	66.99	52.31	40.77	31.78	24.79	19.32	14.97	11.43	8.44	5.85	3.56	1.53*	-0.26*	-1.8	-3.11	-4.21	-5.13	-5.91	-6.56	-7.12	-7.6
$Z_t = 1$	164.76	145.37	125.17	105.32	86.83	70.41	56.4	44.82	35.47	28.02	22.13	17.42	13.6	10.41	7.65	5.21	3.02	1.06*	-0.66*	-2.15	-3.42	-4.48	-5.37	-6.12	-6.76	-7.3	-7.77
$Z_t = (1,t)^t$	158	138.31	118.58	99.73	82.48	67.25	54.23	43.39	34.56	27.45	21.77	17.21	13.49	10.38	7.69	5.31	3.17	1.25*	-0.46*	-1.95	-3.22	-4.29	-5.19	-5.95	-6.6	-7.15	-7.63

Notes: * and in bold indicate non-rejection values at the five percent significance level.

Table 6: FRANCE

Testing the order of integration with the tests of Robinson and white noise disturbances

Z_t/d	0.00	0.05	0.10	0.15	0.20	0.25	0.30	0.35	0.40	0.45	0.50	0.55	0.60	0.65	0.70	0.75	0.80	0.85	0.90	0.95	1.00	1.05	1.10	1.15	1.20	1.25	1.30
$Z_t = 0$	16.99	16.57	16.03	15.34	14.52	13.57	12.48	11.3	10.04	8.74	7.45	6.19	5	3.89	2.88	1.96	1.15*	0.42*	-0.21*	-0.77*	-1.26*	-1.7	-2.08	-2.41	-2.71	-2.97	-3.2
$Z_t = 1$	16.56	15.86	15.12	14.31	13.43	12.47	11.44	10.34	9.19	8.02	6.86	5.71	4.62	3.59	2.64	1.78	0.99*	0.29*	-0.33*	-0.88*	-1.37*	-1.8	-2.18	-2.51	-2.81	-3.07	-3.3
$Z_t = (1,t)^t$	14.15	13.42	12.65	11.82	10.95	10.05	9.12	8.18	7.24	6.32	5.42	4.55	3.71	2.93	2.2	1.51*	0.88*	0.31*	-0.22*	-0.7*	-1.13*	-1.52*	-1.87	-2.18	-2.47	-2.72	-2.96
Testing the order of integration with the tests of Robinson and AR(1) disturbances																											
$Z_t = 0$	148.78	135	118.35	100.36	82.67	66.56	52.77	41.49	32.54	25.55	20.11	15.82	12.36	9.47	6.96	4.71	2.66	0.8*	-0.87*	-2.33	-3.6	-4.69	-5.63	-6.43	-7.13	-7.74	-8.28
$Z_t = 1$	137.23	120.69	104.47	88.61	73.52	59.79	47.91	38.03	30.08	23.79	18.83	14.89	11.67	8.95	6.56	4.39	2.39	0.55*	-1.1*	-2.57	-3.84	-4.93	-5.86	-6.67	-7.36	-7.97	-8.51
$Z_t = (1,t)^t$	88.92	77.68	66.59	56.17	46.79	38.64	31.75	26.03	21.33	17.49	14.32	11.66	9.39	7.38	5.55	3.83	2.21	0.68*	-0.74*	-2.03	-3.19	-4.2	-5.09	-5.87	-6.55	-7.16	-7.7

Notes: * and in bold indicate non-rejection values at the five percent significance level.

Table 7: JAPON

Testing the order of integration with the tests of Robinson and white noise disturbances																										
Z_t/d	0.00	0.05	0.10	0.15	0.20	0.25	0.30	0.35	0.40	0.45	0.50	0.55	0.60	0.65	0.70	0.75	0.80	0.85	0.90	0.95	1.00	1.05	1.10	1.15	1.20	1.25	1.30
$Z_t = 0$	23.49	23.38	23.19	22.9	22.51	21.97	21.27	20.39	19.32	18.08	16.7	15.22	13.7	12.21	10.77	9.43	8.19	7.06	6.03	5.09	4.22	3.43	2.7	2.02	1.4	0.83	0.3
$Z_t = 1$	17.03	16.71	16.37	16	15.59	15.13	14.62	14.06	13.45	12.78	12.06	11.29	10.46	9.6	8.71	7.8	6.87	5.95	5.05	4.17	3.32	2.52	1.78	1.08	0.44	-0.14	-0.67
$Z_t = (1,t)^t$	15.58	15.27	14.94	14.59	14.21	13.8	13.35	12.87	12.35	11.78	11.18	10.53	9.85	9.13	8.38	7.6	6.81	6.01	5.22	4.43	3.66	2.92	2.22	1.55	0.93	0.35	-0.18
Testing the order of integration with the tests of Robinson and AR(1) disturbances																										
$Z_t = 0$	355.81	341.55	317.88	285.49	248.26	210.89	176.56	146.51	120.78	98.97	80.64	65.39	52.87	42.73	34.6	28.14	22.99	18.88	15.56	12.83	10.54	8.55	6.79	5.18	3.69	2.28	0.94
$Z_t = 1$	155.75	148.44	140.36	131.37	121.48	110.83	99.71	88.44	77.41	66.93	57.26	48.55	40.87	34.23	28.55	23.75	19.71	16.29	13.38	10.88	8.67	6.68	4.85	3.14	1.53	0.03	-1.36
$Z_t = (1,t)^t$	123.11	116.73	109.8	102.35	94.49	86.35	78.1	69.93	62.02	54.53	47.57	41.24	35.55	30.52	26.11	22.27	18.94	16.04	13.51	11.28	9.28	7.45	5.75	4.15	2.63	1.19	-0.16

Notes: * and in bold indicate non-rejection values at the five percent significance level.

Figure 1 . Impulse Response of Current Account