MONASH UNIVERSITY

Department of Economics
Discussion Papers
ISSN 1441-5429

The J-Curve: Evidence from Fiji

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No. 01/03

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Abstract

This article provides new evidence on both long-run and short-run determinants of trade balance for Fiji and investigates evidence of *J*-curve adjustment behaviour in the aftermath of a devaluation. We adopt the partial reduced form model of Rose and Yellen (1989) which models the real trade balance directly as a function of the real exchange rate and real domestic and foreign incomes. Cointegration analysis is based on the recently developed Pesaran and Shin (1998) autoregressive distributed lag approach – shown to provide robust results in finite samples – not previously used in the balance-of-trade literature. The long-run elasticities are also estimated using the dynamic ordinary least squares approach of Stock and Watson (1993) and the Fully Modified Ordinary Least Squares (FM-OLS) of Phillips and Hansen (1990). The major results show that there is a long-run relationship between trade balance and its determinants. There is evidence of the J-curve pattern; growth in domestic income affects Fiji's trade balance adversely while foreign income improves it.

Keywords: Fiji, trade balance, cointegration, *J*-curve

Introduction

The balance of payments (BOP) crisis is no stranger to most developing countries, paying frequent visits. The response to the BOP crisis by developing countries is almost exclusively sizeable real devaluations. As a result, in recent years a frequent theme of interest and empirical analysis has hinged on investigating the evidence for the *J*-curve. The *J*-curve is associated with the question of whether devaluation improves the trade balance in the long run; if it does, then the speed of adjustment is crucially important. This question is purely empirical and the consensus developed in the literature is that a real depreciation helps improve the trade balance with a lag of about one-year (Kale, 2001).

The approach taken by many of the empirical studies (for example, Goldstein and Khan, 1985; Bahmani-Oskooee, 1985; Sinha, 2001) have involved the estimation of the structural export and import demand functions to find the price elasticities, while others have followed the non-structural partial reduced form approach introduced by Rose and Yullen (1989).

The empirical literature on the evidence of the *J*-curve is mixed. For instance, Lal and Lowinger (2002a) do not find any evidence of the *J*-curve for Japan. Rose (1990) examined the relationship for a sample of developing countries and found no evidence of the *J*-curve phenomena. Further, Rose and Yellen (1989) examined the short-run relationship between exchange rate and trade balance and found no evidence of the *J*-curve for G-7 countries. Similarly, Wilson and Pat (2001) do not find any evidence of the *J*-curve for Singapore. On the other hand, Demirden and Pastine (1995) uphold the *J*-curve hypothesis for the US. Lal and Lowinger (2002b) find evidence of the *J*-curve for a group of East Asian countries and Kale (2001) finds evidence of the *J*-curve for Turkey.

Using annual time series data for the Fijian economy, this paper attempts to investigate empirically the determinants of trade balance and, in the process, searches for the existence of *J*-curve behaviour in the trade balance. The key reason for undertaking this exercise is that almost all of the present studies are concentrated on testing the *J*-curve hypothesis for large developing economies. The conventional wisdom regarding the validity of any theory is that it gains popularity and greater acceptance if it is empirically tested in countries of various sizes and structures. In this light, Fiji is an exceptional candidate in that it is a small island, open economy with a population of 0.8 million, and is the most socially and economically developed amongst the Pacific Island countries.

Furthermore, Fiji has been struggling over the years to maintain its trade balance at sustainable levels. This owes mainly to its heavy reliance on imports of investment and consumption goods coupled with a narrow range of exports. Faced with BOP crises it had to resort to a 50% devaluation of its currency in the last two decades.

This paper differs from previous studies in that, apart from dealing with a small open economy having a relatively different economic structure, we make a contribution on the methodological front. We draw upon one of the latest advances in econometric time-series modelling with respect to cointegration analysis. In particular, we use the Pesaran and Shin (1998) autoregressive distributed lag (ARDL) approach to cointegration – also known as the bounds test procedure - to investigate the existence of a long-run relationship. Pesaran and Shin (1998) show that, under the ARDL framework, the OLS estimators of the short run parameters are $\sqrt{T}$-consistent, and the ARDL based estimators of the long run coefficients are super-consistent in small sample sizes. This methodology has not been previously used in the trade balance literature. We go even further: for the first time in the trade balance literature we utilise Hansen (1992) test for stability of the cointegration relationship. There is no reason to believe a priori, as is the case in the trade balance literature, that the relative importance of factors influencing trade balance has remained unchanged.

The aims of this paper are accomplished in four steps. First, the time series univariate properties are diagnosed by a number of unit root tests. The commonly used Dickey and Fuller (1979, 1981) test are used, as is a tradition in applied time series analysis. However, given the recent criticisms it has received for having limited power in small samples (Blough, 1992; Inder, 1994), and given our small sample size, we also apply the Phillip-Perron (1988) and the Kwiatkowski *et al.* (1992) tests. Second, the ARDL approach is used to investigate the existence of any long-run relationships between trade balance, real effective exchange rate, foreign income and domestic income. Third, the long-run estimates are obtained by using three different methods: the ARDL, the Dynamic Ordinary Least Squares (DOLS) of Stock and Watson (1993), and the Fully Modified Ordinary Least Squares (FM-OLS) of Phillips and Hansen (1990). The reasons for using these methods are twofold: first, they provide more efficient results in small samples and second, they provide a good basis for the comparison of the robustness of results. The short-run estimates are derived using the ARDL approach. Fourth, Hansen stability tests are conducted. Finally, the generalised variance decomposition analysis and the generalised impulse response functions are used to investigate the response of trade balance to real effective exchange rate shocks. The idea here is to test the *J*-curve hypothesis.

The paper proceeds as follows. The next section presents the trade balance model to be estimated. This is followed by a description of the methodologies used in this study. The penultimate section contains the empirical findings. In the last section, the conclusions of the analysis are summarised.

Model and Methodology

We utilise the non-structural partial reduced form model of Rose and Yellen (1989), which is as follows:

$$B = f(REER, FY, GNI) \tag{1}$$

Instead of using net exports as the dependent variable, we use the ratio of imports to exports (ME). This enables us to transform the variables in logarithmic form and allows the coefficients to be interpreted as elasticities. The trade balance model for Fiji takes the following form:

$$lnME_t = \alpha + \beta_1 lnREER_t + \beta_2 lnFY_t + \beta_3 lnGNI_t + \varepsilon_t \tag{2}$$

where $lnME_t$ is the logarithm of the real imports to real exports ratio; $lnREER_t$ is the logarithm of the trade weighted real effective exchange rate, defined as the number of units of domestic currency per unit of foreign currency; $lnFY_t$ is the logarithm of the weighted average of trading partners' real income which captures the trading partners demand conditions and $lnGNI_t$ is the logarithm of real domestic income; α is a constant; ε_t is a error term; and β_1, β_2 and β_3 are parameters to be estimated.

According to the *J*-curve hypothesis, an increase in real effective exchange rate initially reduces the demand for the home country's exports but increases its demand for imports. This initially leads to a deterioration of the BOT due to the belief that imports in local currency increase more than the initial increase in exports after a change in price. However, as export and import volumes adjust to price changes over time, the BOT improves. Hence, it is expected that $\beta_1 < 0$. The signs associated with β_2 and β_3 are purely empirical. For instance, an increase in the economic activity of a trading partner country not only boosts its demand for imports from Fiji but also its supply of exports to Fiji; hence, β_2 could be a negative or a positive depending on whether demand side factors dominate supply side factors or *vice versa*.

The time-series data adopted for this study are annual and cover the period from 1970 to 1999. The data series are sourced from the IMF International Financial Statistics, the Fiji Bureau of Statistics' Current Economic Statistics and the Reserve Bank of Fiji Quarterly Reviews. The ARDL approach is used to test for the existence of any long-run relationships, while the ARDL, DOLS and the FM-OLS are used to estimate the long-run elasticities. In what follows, we describe these methodologies.

ARDL Approach: The augmented autoregressive distributed lag (ARDL) $(p, q_1, q_2, \ldots, q_k)$ model can be written as follows (Pesaran and Pesaran, 1997: 397-399; Pesaran *et al.* 2001):

$$\Omega(L,p)y_t = \alpha_0 + \sum_{i=1}^{k} \beta_i(L,q_i)x_{it} + \delta' w_t + \mu_t \tag{3}$$

where

$$\Omega(L,p) = 1 - \Omega_1\delta_1 L^1 - \Omega_2\delta_2 L^2 - \ldots - \Omega_p L^p, \tag{4}$$

$$\beta_i(L,q_i) = \beta_{io} + \beta_{i1}L + \beta_{i2}L^2 + \ldots + \beta_{iqi}L^{qi},\ i = 1,2,\ldots,\ k, \tag{5}$$

Here, y_t is the dependent variable; α_o is a constant; L is a lag operator such that $Ly_t = y_{t-1}$; and w_t is a $s \times 1$ vector of deterministic variables such as seasonal dummies, time trends, or exogenous variables with fixed lags. The x_{it} in equation (3) is the i independent variable where i=1,2,...,k. In the long-run, we have $y_t = y_{t-1} = \ldots = y_{t-p}$; $x_{it} = x_{i,t-1} = \ldots = x_{i,t-q}$ where $x_{i,t-q}$ denotes the q^{th} lag of the i^{th} variable.

The long-run equation with respect to the constant term can be written as follows:

$$y = \alpha_o + \sum_{i=1}^{k} \beta_i x_i + \delta' w_t + v_t \qquad \Omega = \frac{\alpha_0}{\Omega(1,p)} \tag{6}$$

The long-run coefficient for a response of y_t to a unit change in x_{it} are estimated by:

$$\beta_i = \frac{\hat{\beta}_i(1,\hat{q}_i)}{\hat{\Omega}(1,\hat{p})} = \frac{\hat{\beta}_{i0} + \hat{\beta}_{i1} + \ldots + \hat{\beta}_{i\hat{q}}}{1 - \Omega_1 - \hat{\Omega}_2 - \ldots - \hat{\Omega}_{\hat{p}}},\ i = 1,2,\ldots,k \tag{7}$$

Here, $\hat{p}$ and $\hat{q}_i$, $i = 1,2,\ldots,k$ are the selected (estimated) values of p and q_i, $i = 1,2,\ldots k$. Similarly, the long-run coefficients associated with the deterministic/exogenous variables with fixed lags are estimated by

$$\delta' = \frac{\hat{\delta}(\hat{p},\hat{q}_1,\hat{q}_2,\ldots,\hat{q}_k)}{1 - \hat{\Omega}_1 - \hat{\Omega}_2 - \ldots - \hat{\Omega}_p}, \tag{8}$$

Here, $\hat{\delta}(\hat{p},\hat{q}_1,\hat{q}_2,\ldots,\hat{q}_k)$ denotes the ordinary least squares estimate of δ in equation (2) - the selected ARDL model. The error correction (EC) representation of the ARDL$(\hat{p},\hat{q}_1.\hat{q}_{2,\ldots,}\hat{q}_k)$ model can be obtained by writing equation (2) in terms of the lagged levels and the first differences of $y_t, x_{1t}, x_{2t,\ldots,} x_{kt}$ and w_t:

$$\Delta y_t = \Delta\alpha_0 - \sum_{j=1}^{\hat{p}-1} \Omega_j^{*} \Delta y_{t-j} + \sum_{i=1}^{k} \beta_{i0} \Delta x_{it} - \sum_{i=1}^{k} \sum_{j=1}^{\hat{q}_i-1} \beta_{ij}^{*} \Delta x_{i,t-j} + \delta' \Delta w_t - \Omega(1,\hat{p}) ECM_{t-1} + \mu_t \tag{9}$$

Here, ECM_t is the correction term defined by

$$ECM_t = y_t - \hat{\alpha} - \sum_{i=1}^{k} \hat{\beta}_i x_{it} - \delta' w_t \tag{10}$$

and Δ is the first difference operator; Ω_j^{*}, β_{ij}^{*} and δ' are the coefficients relating to the short-run dynamics of the model's convergence to equilibrium while $\Omega(1,\hat{p})$ measures the speed of adjustment.

The bounds testing procedure involves two stages. The first stage is to establish the existence of a long-run relationship. Once a long-run relationship has been established, a two step procedure is used in estimating the long-run relationship. An initial investigation of the existence of a long-run relationship predicted by theory among the variables in question (see equation 11 below) is preceded by an estimation of the short-run and long-run parameters using equation (2). Suppose that with respect to equation 2, theory predicts that there is a long-run relationship among $LME_t, LREER_t, LFY_t$ and $LGNI_t$. Without having any prior information about the direction of the long-run relationship among the variables, the following unrestricted error correction (EC) regressions are estimated, taking each of the variables in turn as a dependent variable:

$$
\begin{aligned}
\Delta ln ME_t = a_{0ME} + \sum_{i=1}^{n} b_{iME} \Delta ln ME_{t-i} + \sum_{i=0}^{n} c_{iME} \Delta ln REER_{t-i} \\
+ \sum_{i=0}^{n} d_{iME} \Delta ln FY_{t-i} + \sum_{i=0}^{n} e_{iME} \Delta ln GNI_{t-i} + \lambda_{1ME} ln ME_{t-i} \\
+ \lambda_{2ME} ln REER_{t-i} + \lambda_{3ME} ln FY_{t-i} + \lambda_{5ME} ln GNI_{t-i} + \varepsilon_{1t}
\end{aligned} \qquad (11a)
$$

$$
\begin{aligned}
\Delta ln REER_t = a_{0REER} + \sum_{i=1}^{n} b_{iREER} \Delta ln REER_{t-i} + \sum_{i=0}^{n} c_{iREER} \Delta ln ME_{t-i} \\
+ \sum_{i=0}^{n} d_{iREER} \Delta ln FY_{t-i} + \sum_{i=0}^{n} e_{iREER} \Delta ln GNI_{t-i} + \lambda_{1REER} \Delta ln REER_{t-i} \\
+ \lambda_{2REER} \Delta ln ME_{t-i} + \lambda_{3REER} ln FY_{t-i} + + \lambda_{5REER} ln GNI_{t-i} + \varepsilon_{2t}
\end{aligned} \qquad (11b)
$$

$$
\begin{aligned}
\Delta ln FY_t = a_{0FY} + \sum_{i=1}^{n} b_{iFY} \Delta ln FY_{t-i} + \sum_{i=0}^{n} c_{iFY} \Delta ln REER_{t-i} \\
+ \sum_{i=0}^{n} d_{iFY} \Delta ln ME_{t-i} + \sum_{i=0}^{n} e_{iFY} \Delta ln GNI_{t-i} + \lambda_{1FY} \Delta ln REER_{t-i} \\
+ \lambda_{2FY} \Delta ln ME_{t-i} + \lambda_{3FY} ln FY_{t-i} + \lambda_{5FY} ln GNI_{t-i} + \varepsilon_{3t}
\end{aligned} \qquad (11c)
$$

$$
\begin{aligned}
\Delta LGNI_t = a_{0GNI} + \sum_{i=1}^{n} b_{iGNI} \Delta ln GNI_{t-i} + \sum_{i=0}^{n} c_{iGNI} \Delta ln REER_{t-i} \\
+ \sum_{i=0}^{n} d_{iGNI} \Delta ln ME_{t-i} + \sum_{i=0}^{n} e_{iGNI} \Delta ln FY_{t-i} + \sum_{i=0}^{n} f_{iGNI} \Delta ln FR_{t-i} + \lambda_{1GNI} ln REER_{t-i} \\
+ \lambda_{2GNI} ln ME_{t-i} + \lambda_{3GNI} ln FR_{t-i} + \lambda_{4GNI} ln FY_{t-i} + \lambda_{5GNI} ln GNI_{t-i} + \varepsilon_{3t}
\end{aligned} \qquad (11d)
$$

When a long-run relationship exists, the F test indicates which variable should be normalised. The null hypothesis for no cointegration amongst the variables in Equation 11a is $(H_0 : \lambda_{1ME} = \lambda_{2ME} = \lambda_{3ME} = \lambda_{4ME} = 0)$ denoted by $F_{ME}(ME|REER, FY, GNI)$ against the alternative $(H_1 : \lambda_{1ME} \neq \lambda_{2ME} \neq \lambda_{3ME} \neq \lambda_{4ME} \neq 0)$. Similarly, the null hypothesis for testing the 'nonexistence of a long run relationship' in equation (11b) is denoted by $F_{REER}(REER|ME, FY, GNI)$; for equation (11c) the F test for testing the null hypothesis is denoted by $F_{FY}(FY|ME, REER, GNI)$; and for equation (11d) the F test is denoted by $F_{GNI}(GNI|ME, REER, FY)$.

The F test has a non-standard distribution which depends upon; (i) whether variables included in the ARDL model are $I(0)$ or $I(1)$, (ii) the number of regressors and (iii) whether the ARDL model contains an intercept and/or a trend. Two sets of critical values are reported in Pesaran and Pesaran (1997) (see also Pesaran *et al.* 2001). The two sets of critical values provide critical value bounds for all classification of the regressors into purely $I(1)$, purely $I(0)$ or mutually cointegrated.

If the computed F statistics falls outside the critical bounds, a conclusive decision can be made regarding cointegration without knowing the order of integration of the regressors. For instance, if the empirical analysis shows that the estimated $F_{ME}(.)$ is higher than the upper bound of the critical values then the null hypothesis of no cointegration is rejected. Once a long-run relationship has been established, in the second stage, a further two step procedure to estimate the model is carried out. First the orders of the lags in the ARDL model are selected using an appropriate lag selection criteria such as the Schwartz Bayesian Criterion (SBC) and in the second step the selected model is estimated by the ordinary least squares technique.

Dynamic DOLS: The procedure advocated by Stock and Watson (1990) involves estimation of long run equilibria via dynamic OLS (DOLS), which corrects for potential simultaneity bias among regressors. It resembles the ideas inherent in Hansen (1988), Phillips and Loretan (1991), Phillips and Hansen (1990), Saikkonen (1991), and Park (1992). DOLS entails regressing one of the $I(1)$ variables on other $I(1)$ variables, the $I(0)$ variables, and lags and leads of the first difference of the $I(1)$ variables. The essence of incorporating the first difference variables and the associated lags and leads is to obviate simultaneity bias and small sample bias inherent among regressors. Standard hypothesis testing can be undertaken using robust standard errors derived via the procedure recommended by Newey and West (1987).

The DOLS is based on an alternative representation of the system which assumes the following particularly a priori normalisation, which can be obtained in any system with cointgerating vectors:

$$\Delta X_t^1 = \kappa_t^1$$

$$X_t^2 = \Phi_0 + \Phi X_t^1 + \kappa_t^1 \qquad (12)$$

where $X_t' = \left[X_t^{1'} \middle| X_t^{2'}\right]$, the dimensions of X_t^1 and X_t^2 being $(p-r)\times 1$ and $(r \times 1)$, respectively. The error processes are deemed stationary and by incorporating both leads and lags of ΔX_t^1 in equation (12) and estimating the normalised cointegrating vectors, Φ, by OLS, one can obtain an estimator asymptotically equivalent to MLE.

Fully Modified OLS (FMOLS): The procedure, developed by Phillips and Hansen (1990), has two direct advantages. Apart from correcting for where $X_t' = [X_t^{1'} | X_t^{2'}]$, the dimensions of X_t^1 and X_t^2 being $(p-r)\times 1$ and $(r \times 1)$, respectively. The error processes are deemed stationary and by incorporating both leads and lags of ΔX_t^1 in equation (12) and estimating the normalised cointegrating vectors, Φ, by OLS, one can obtain an estimator asymptotically equivalent to MLE.

Fully Modified OLS (FMOLS): The procedure, developed by Phillips and Hansen (1990), has two direct advantages. Apart from correcting for endogeneity and serial correlation effect it also asymptotically eliminates the sample bias. There are two conditions considered essentially for the appropriateness of the FMOLS. First, there is only one integrating vector. Second the explanatory variables are not cointegrated among themselves. Assuming these provisions are met, the econometric model is of the following form:

$$y_t = \sigma_0 + \sigma_1' X_t + \mu_t, \qquad t = 1,2,\ldots,n$$

where y_t is an $I(1)$ variable and X_t is a $(k \times 1)$ vector of $I(1)$ regressors, which are not cointegrated among themselves. By assumption, X_t has the following first difference stationary process:

$$\Delta X_t = \eta + \lambda_t, \qquad t = 2,3,\ldots,n$$

where η is a $k \times 1$ vector of drift parameters, λ_t is a $k \times 1$ vector of $I(0)$ variables. It is also assumed that $\varpi_t = (\mu_t, \lambda_t')'$ is strictly stationary with zero mean and a finite positive-definite covariance matrix, Σ.

The ARDL approach does not require knowledge about the order of integration of variables in searching for cointegration relationships; however, the DOLS and FM-OLS, for the purpose of estimating long-run relationships, require knowledge of the integration properties. To accomplice this, the variables in the long run equilibrium model are tested for unit roots or for non-stationarity.[1] The respective unit root tests are as follows.

The widely used unit root tests is the Augmented Dickey-Fuller (ADF) (1979) and the Phillips – Perron (1988) test.[2] The ADF test is based on the following regression equation:

$$\Delta x_t = a_0 + \lambda T + \phi x_{t-1} + \sum_{i=1}^{p-1} \gamma_i \Delta x_{t-1} + e_t$$

where x_t is the variable tested for unit root; Δ is the first difference operator; α is the constant; T is the time trend variable; and p is the number of lag included to avoid the problem of autocorrelation in the residuals. The lag length in the ADF regression is selected based on the minimum Schwarz Bayesian criterion. The null hypothesis in the ADF tests is that the series (which should be in level form) is non-stationary, i.e., it contains unit root. To reject the null, the calculated test value has to be greater than the critical value. The critical values are calculated from MacKinnon (1991).

The Phillips and Perron (1988) is an alternative to the ADF test. It controls for serial correlation when testing for unit root and is based on the non-augmented DF-test equation. The key focus of this method is on modifying the t-ratio so that serial correlation does not affect the asymptotic distribution of the test statistic (Eviews 4.1, 2002).

The KPSS (1992) test for unit root differs from the ADF and the PP test in that the series Y_t is assumed to be (trend-) stationary under the null. Put differently, KPSS test reverses the null and the alternative hypothesis. The

KPSS statistic is based on the residuals from the OLS regression which takes the following form:

$$Y_t = \alpha_t + b_t + \varepsilon_t$$

where t is a linear deterministic trend, ε_t is a stationary error, and b_t is a random walk; $b_t = b_{t-1} + \mu_t$, where μ_t are i.i.d. (0, σ^2_μ). The initial value of b_0 is treated as fixed and is interpreted as an intercept. The test is conducted by first regression Y_t on a constant and a trend (t). This allows to obtain the residuals. The KPSS statistic is defined as:

$$\eta = T^2 \sum S_t^2 / S^2(k),$$

where $S_t = \sum^t e_i (i = 1,2,\ldots\ldots,t)$ and e_i is the partial sum of the residuals $S^2(k)$ is a consistent non-parametric estimate of the disturbance variance and T is the sample size.

Empirical Results

Unit root tests

Table 1 reports the unit root tests. The ADF and PP statistics for the import to export ratio, REER, domestic income and foreign income do not exceed the critical values (in absolute terms). We therefore could not find any significant evidence that $[ME_t, REER_t, GNI_t, FY_t]$ are not integrated of order one or $I(1)$.

INSERT TABLE 1

On the other hand, the KPSS test statistic that tests the null hypothesis that a particular variable is mean stationary exceeds the critical values; consistent with our findings from the ADF and PP tests that all variables in our model have unit roots and are clearly nonstationary in levels. However, when all these variables are differenced once and subjected to the ADF and PP tests, we find that the test statistics exceed the critical values. On the other hand the KPSS tests of the first difference variables proffer test statistics smaller than the critical values. This leads us to the conclusion that all variables $[ME_t, REER_t, GNI_t, FY_t]$ are stationary in their first differences.

Cointegration test

In the first step of the ARDL analysis we tested for the presence of long-run relationships in equations (11a), (11b), (11c) and (11d). As we use annual data, the maximum number of lags in the ARDL was set equal to 2. The calculated F-statistics are reported in Table 2. For equation 11a, $F_{ME}(.) = 6.705$ is higher than the upper bound critical value of 5.615 at the 1 per cent level. Thus, the null hypothesis of no cointegration cannot be accepted; there is a cointegration relationship amongst the variables in each of the models.

INSERT TABLE 2

Once we established that a long-run cointegration relationship existed, equation (2) was estimated using the following ARDL (m, n, p, q) specification:

$$ln\ ME_t = \alpha_0 + \sum_{i=1}^{m} \alpha_1\, ln\ ME_{t-i} + \sum_{i=0}^{n} \alpha_2\, ln\ REER_{t-i} +$$
$$+ \sum_{i=0}^{p} \alpha_3\, ln\ FY_{t-i} + \sum_{i=0}^{q} \alpha_4\, ln\ GNI_{t-i} + \mu_t$$
(13)

For each model a maximum of 2 lags was used such that i_{max}=2. The estimated model presented here is based on the SBC. Results of the long run model estimated by using the ARDL, together with estimates from the FMOLS and DOLS are presented in Table 4. The robustness of the long run results is verified partly by the fact that all three methods provide parallel results.

Apart from a robust long-run relationship, our short-run error correction model is statistically well behaved. The error correction term EC_{t-1}, which measures the speed of adjustment to restore equilibrium in the dynamic model, has a negative sign and is statistically significant at the 1% level ensuring that the series is non-explosive and that long-run equilibrium is attainable. The coefficient of –0.71 implies that a deviation from the long-run ME ratio during this period is corrected by about 71 percent in the next period– an indication that, following a shock, convergence to equilibrium is swift.

We also applied a number of diagnostic tests to the error correction model (Table 3). There is no evidence of autocorrelation in the disturbance of the error term. The ARCH tests suggest the errors are homoskedastic and independent of the regressors. The model passes the Jarque-Bera normality tests suggesting that the errors are normally distributed. The RESET test indicates that the model is correctly specified, while the F-forecast test indicates the predictive power/accuracy of the model. Finally, the adjusted R-squared of the model, 0.67, is reasonable – 67% of the variations in the ME ratio is explained by the regressors.

INSERT TABLE 3

INSERT TABLE 4

The long-run elasticity of the ME ratio with respect to the real effective exchange rate is negative, implying that a devaluation of the real exchange rate will lead to a reduction in the ratio of imports to exports. Put differently, it indicates an improvement in the trade balance. The elasticity ranges between –0.1 to –0.2, implying an inelastic response of the ME ratio with respect to changes in the real effective exchange rate. This result is, however, statistically insignificant.

There is no consensus in the literature regarding the direction of the relationship between the trade balance and domestic and foreign income. An increase in the income of a country or that of its major trading partners, for instance, can induce an increase in the supply of all goods including tradeable goods of a certain country or that of its trading partners. It is also true that an increase in income in a country or its major trading partners can increase demand for all goods including tradeable goods. It follows, then, that this ambiguity, whether demand side factors dominate supply side factors or vice versa, is purely an empirical issue. Our long-run results indicate that an increase in income in Fiji increases the import to export ratio, implying a deterioration of the trade balance. The domestic income elasticity is elastic, ranging between 1.2 to 1.5.

On the other hand, an increase in foreign income exerts a negative influence on the trade balance – a fall in the import to export ratio. This implies an improvement in Fiji's trade balance in the long-run. The response of the trade balance to foreign income is fairly inelastic, ranging between –0.3 to –0.4. The results of the income effects seem reasonable and plausible as follows: Firstly, Fiji's import demand is mostly made up of consumption goods (40%), and investment and energy goods (50%). Therefore the likelihood of demand for these goods increasing with higher domestic income is extremely high. Secondly, goods exported by Fiji are mostly primary commodities (57%) and manufactured garments (30%). Demand for these exports does not seem to be (foreign) income elastic.

Constancy of Cointegration Space

One problem with time series regression models is that the estimated parameters may change over time. Unstable parameters can result in model misspecification and, if left undetected, have the potential to bias the results. To account for this, here, we examine whether the estimated elasticities are stable over time. To do this we use the parameter non-constancy tests for $I(1)$ processes advocated by Hansen (1992). Hansen (1992) proposes three tests – $SupF$, $MeanF$, and L_C – which all have the same null hypothesis that there is no structural change but differ in their choice of alternative hypothesis. The $SupF$ test is predicated on ideas inherent in the classical Chow F-tests. The alternative hypothesis is a sudden shift in regime at an unknown point in time, and amounts to calculating the Chow F-statistic. This test statistic takes the following form: $SupF = SupF_{t/T}$, where $F_{t/T}$ is the F-test statistic. To perform the $SupF$ test requires truncation of the sample size T. We follow the approach in Hansen (1992) and use the subset $[0.15T, 0.85T]$.

The *MeanF* test is appropriate when the question under investigation is whether or not the specified model is a good model that captures a stable relationship (Hansen, 1992). It is computed as an average of the $F_{t/T}$. Finally, the L_C statistic is recommended if the likelihood of parameter variation is relatively constant throughout the sample. The test results and their probability values are reported in Table 5. They indicate parameter stability, since the probability values for each test are significant at the 5 per cent level. This indicates that the structure of the parameters have not diverged abnormally over the period of the analysis.

Variance decomposition

The essence of a variance decomposition analysis is its ability to provide information about the relative importance of the random innovations. Specifically, it provides information on the percentage of variation in the forecast error of a variable explained by its own innovation and the proportion explained by innovations in other variables. Table 6 summarises the results of the variance decomposition on the effects of trade balance, real effective exchange rates, foreign income and domestic income on the trade balance.

INSERT TABLE 6

The analysis of variance decomposition suggest that a significant percentage of the variability in trade balance - on average 65.3% of the variation in the forecast error for trade balance - can be explained by its own innovations which, however, decline overtime. The results demonstrate that current performance of trade balance is contingent largely upon past performances. Some 12.7% and 21.6% of the forecast error variance for trade balance is explained by foreign income and domestic income, respectively. The empirical results further suggest that only a small proportion of variability in trade balance could be attributed to innovations in the real effective exchange rates.

Impulse response function

An alternative way of obtaining information regarding the relationships among the variables included in the variance decomposition analysis is via the generalized impulse response functions. The generalized impulse response functions reveal insights into the dynamic relationships in existence, as they portray the response of a variable to an unexpected shock in another variable over a certain time horizon.

INSERT FIGURE I

The impulse response function reveals that an increase in the REER (or devaluation of Fiji's currency) leads to an initial rise in the ratio of imports to exports – a deterioration of the trade balance – for the first 2 years. This deterioration is followed by an improvement in Fiji's trade balance as indicated by a fall in the ratio of imports to exports. This result indicates the presence of a J-curve pattern of trade balance for Fiji. The three devaluations of the Fiji dollar since 1987 have been undertaken with the aim of increasing Fiji's export competitiveness. The results here provide evidence that this aim was achieved in the years following the first 2 years from the time when the devaluation policy was taken.

Conclusion

This paper uses a reduced form trade balance model to investigate the existence of the J-curve phenomenon and, at the same time, to highlight the determinants of the trade balance of Fiji, using modern econometric techniques. Using the ARDL technique, we show a long run relationship between the import to export ratio, the real effective exchange rate, and foreign and domestic incomes. Upon estimating the trade balance model using the three different approaches (ARDL, DOLS, and FMOLS), it was clear that domestic income is the most import determinant of the trade balance, followed by foreign income.

Four important results emerge from our analysis. First, a real devaluation of the Fiji dollar leads to an improvement in the trade balance in the long-run. Second, an increase in domestic income adversely affects the trade balance in the long-run suggesting that the expansionary effects of growth on imports outweighs the positive impact of exportable surplus. Third, Fiji's trade balance is likely to improve from growth in foreign income, although the response is relatively inelastic. Fourth, the impulse response analysis gives evidence pointing to a deterioration in the Fijian trade balance due to a shock in the real effective exchange rate for the first 2 years; thereafter, the trade balance starts improving. This is testimony to the existence of a J-curve effect.

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Table 1: Unit root test results: ADF, PP and KPSS

Variables	ADF stat [LL]	CV	PP stat [BW]	CV	KPSS stat [BW]	CV
ln ME	-2.768 [0]	-3.568	-3.112 [6]	-3.568	0.165 [4]	0.146
ln REER	-2.185 [0]	-3.568	-2.185 [0]	-3.568	0.188 [3]	0.146
ln FY	-2.258 [0]	-3.574	-2.257 [2]	-3.574	0.194 [4]	0.146
ln GNI	-4.058 [8]	-3.633	-3.288 [8]	-3.568	0.196 [3]	0.146
Δln ME	-7.188 [0]	-3.574	-6.932 [3]	-3.574	0.103 [2]	0.146
Δln REE	-5.137 [0]	-3.574	-5.143 [2]	-3.574	0.067 [2]	0.146
Δln FY	-5.217 [0]	-3.581	-5.216 [1]	-3.581	0.084 [0]	0.146
Δln GNI	-4.178 [0]	-3.574	-4.178 [0]	-3.574	0.112 [0]	0.146

Note: LL is Lag Lenth; CV is Critical values at 5% level; and BW is the Bandwidth.

Table 2: F-statistics for Cointegration Relationship

Critical value bounds of the F statistic						
	95% level		97.5% level		99% level	
k	$I(0)$	$I(1)$	$I(0)$	$I(1)$	$I(0)$	$I(1)$
3	3.219	4.378	3.727	4.898	4.385	5.615
Calculated F statistics						
$F_{ME}(ME \mid REER, FY, GNI) = 6.7053$						
$F_{REER}(REER \mid ME, FY, GNI) = 3.9000$						
$F_{FY}(FY \mid ME, REER, GNI) = 0.6529$						
$F_{GNI}(GNI \mid ME, REER, FY) = 4.2289$						

Notes: The critical value bounds are from Table F in Pesaran and Pesaran, (1997: 484). k is the number of regressors.

Table 3: Fiji's Trade Balance in the Long Run, 1970-2000

ARDL Approach - ARDL (0,1,0,1,1) selected based on SBC			
Regressors	Coefficient	Standard error	t-statistics
Constant	-0.3607	4.3689	-0.8257
$LREER_t$	-0.1055	0.3810	-0.2769
LFY_t	-0.4085***	0.1581	-2.5837
$LGNI_t$	1.2124***	0.2155	5.6251
Fully Modified Phillips-Hansen Estimates, 1970-2000			
Constant	-5.5685*	3.2844	-1.7142
$LREER_t$	-0.2137	0.2948	-0.7248
LFY_t	-0.3093***	0.1164	-2.6565
$LGNI_t$	1.4163***	0.1442	9.8243
DOLS Estimates, 1970-2000			
Constant	-1.7135	2.4235	-0.7071
$LREER_t$	-0.1261	0.2215	-0.5693
LFY_t	-0.4470***	0.0985	4.5380
$LGNI_t$	1.5011***	0.1648	9.1082

$R^2 = 0.9670$

$\bar{R}^2 = 0.9428$

$SER = 0.0434$

Notes: *(***) indicates statistical significance at the 10% and 1% levels respectively.

Table 4: Fiji's Trade Balance in the Short Run: Estimates from Error Correction Model

Regressors	Coefficient	Standard error	t-statistics
Constant	-0.2570	3.1338	-0.8202
$LREER_t$	-0.6444***	0.2778	-2.3191
LFY_t	0.1158	0.1542	0.7512
$LGNI_t$	0.1581	0.3300	0.4790
$ECM(-1)$	-0.7125***	0.1296	-5.4990

$R^2 = 0.7680$	$\chi^2_{NORM}(2) = 1.1966$	$\chi^2_{RESET}(2) = 4.7272$
$\bar{R}^2 = 0.6752$	$\chi^2_{AUTO}(2) = 0.1029$	
$\sigma = 0.0689$	$\chi^2_{WHITE}(10) = 12.0771$	
$\chi^2_{NORM}(2) = 0.4648$	$F_{FORCAST}(1990) = 0.2197$	

Note: *** indicates statistical significance at the 1% level.

Table 5: Hansen test for parameter stability

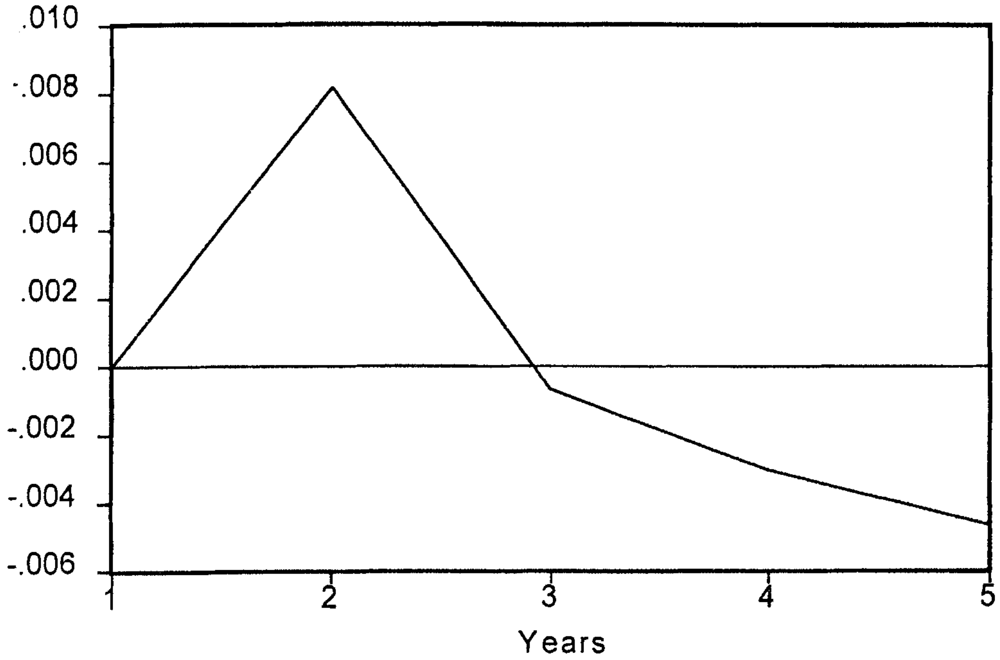
Tests	Test statistic	Probability value
L_C	0.2509	>0.20
MeanF	1.2606	>0.20
SupF	4.5611	>0.20

Note: The test program is available from http//www.ssc.wisc.edu/bhansen/

Table 6: Variance decomposition on the effects of trade balance, real effective exchange rates, foreign reserves, foreign income and domestic income on the trade balance for Fiji

Period	Trade balance	REER	Foreign income	Domestic income
1	100.000	0.0000	0.000	0.0000
2	62.362	0.314	13.266	24.058
3	64.341	0.387	12.830	22.441
4	61.895	0.454	13.922	23.728
5	62.085	0.469	13.904	23.541
6	61.814	0.479	14.032	23.675
7	61.832	0.481	14.034	23.653
8	61.800	0.482	14.050	23.668
9	61.802	0.482	14.051	23.665
10	61.798	0.483	14.052	23.667
11	61.798	0.483	14.053	23.667
12	61.797	0.483	14.053	23.667

Figure 1: Impulse response function: Effect of one standard deviation innovation in the real effective exchange rate on trade balance (import to export ratio) in Fiji



Endnotes

[1] A time series process is said to be stationary it does not follow a trend but follows a random path such that the mean of the process is constant over time. More so, a non-stationary series can be made stationary by a commonly used method – which is to difference the series d times. A non-stationary series which can be transformed to a stationary series by differencing d times is said to be integrated of order d, denoted $x_t \sim I(d)$ (Engle and Granger, 1987).

[2] See Said and Dickey (1984) and Phillips and Perron (1988).

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