# The Determinants of Immigration from Fiji to New Zealand: An Empirical Reassessment Using the Bounds Testing Approach[1]

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## ABSTRACT

This article re-examines Gani's (1998) findings on the determinants of migrant flows from Fiji to New Zealand by employing the bounds testing procedure to cointegration, within an autoregressive distributive lag framework. The main findings are that in the long run all variables are statistically insignificant, although correctly signed with the exception of the unemployment differential. In the short run, in sharp contrast to Gani's (1998) findings, political instability is consistently the most important determinant of migration flows while the standard of living and real wage differentials are statistically insignificant across all specifications.

## INTRODUCTION

New Zealand is the third major destination for immigrants from Fiji. In 1997, of the total number of immigrants leaving Fiji, 33 per cent went to the United States, 31 per cent went to Australia, 21 per cent went to New Zealand, and 10 per cent went to Canada (Gani, 2000: 96-97; Mohanty, 2002a: 7). Prior to 1987 only about 5,500 Fijians had migrated to New Zealand. The pace of immigration from Fiji to New Zealand, however, has intensified following the first Fiji coups in 1987. In the 1996 census there were 19,000 Fiji-born people living in New Zealand, of whom 11,880 (64%) had been there for less than nine years (Mohanty, 2002a: 8).

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The increase in the level of migration from Fiji to New Zealand following the first coup in 1987 has mirrored a more general trend in the number of immigrants leaving Fiji. The official statistics indicate that about 75,000 Fiji citizens have immigrated with an annual average of more than 5,000 people between 1987 and 2001. The rate of immigration surged following the 2000 coup. In 2000 and 2001 alone, the number of immigrants was 5,275 and 6,316 respectively (Kumar and Prasad, 2002: 11). These are significant numbers given that Fiji has a population of 775,000.

Previous studies of migration from Fiji to New Zealand include both non-econometric studies (Connell, 1985, 1987; Bedford and Levick, 1988) and econometric studies (Gani and Ward, 1995; Gani, 1998). The latter two studies are a subset of a growing literature which has examined the determinants of migration from developing to developed countries. Other studies include Devoretz and Maki (1983) and Akbar and Devoretz (1993) who consider the determinants of migration from developing countries to Canada. Huang (1987) and MacPhee and Hassan (1990) examine the determinants of migration flows to the United States and Lam (2002) explores the reasons for migration from Hong Kong following the end of British rule.

Of the previous econometric studies of the determinants of migration from Fiji to New Zealand, Gani and Ward (1995) examine the determinants of migration of skilled professionals using panel data for 19 occupational groups between 1987 and 1990. Their main finding was that the number of professional migrants from Fiji was positively related to real income in New Zealand and political instability in Fiji. Gani (1998) examines the determinants of immigration from Fiji to New Zealand within a human capital framework using time series data from 1970 to 1994. He uses the error correction framework with the Engle-Granger (1987) approach to cointegration. His main finding was that unemployment and wage differentials were the major determinants of migration from Fiji to New Zealand, while living standard differentials and political instability in Fiji were statistically insignificant.

The objective of this article is to re-examine Gani's (1998) findings on the determinants of migrant flows from Fiji to New Zealand using time series data from 1972 to 2001. This study differs from Gani's (1998) in the following respects.

First, we use a slightly longer time period, which has the advantage of encompassing the 2000 coup and, therefore, takes account of the effect of continuing political instability in Fiji on migration decisions after Gani's (1998) study concluded.

Second, whereas Gani (1998) employs the Engle-Granger (1987) approach to cointegration we use the bounds testing procedure to cointegration, within an autoregressive distributive lag (ARDL) framework, developed by Pesaran and others (Pesaran and Pesaran, 1997; Pesaran and Shin, 1999; Pesaran, et al., 2001). Given the relatively small sample size in both the current study and that of Gani (1998), the bounds testing approach to cointegration is preferable to the Engle-Granger (1987) method of cointegration. It is well known that estimates using the Engle-Granger (1987) method of cointegration are not robust for small sample sizes (see, e.g. Mah, 2000; Tang and Nair, 2002). However, Pesaran and Shin (1999) show that with the ARDL framework, the ordinary least squares (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. Previous studies have applied the bounds test to relatively small sample sizes with as few as 20 observations in considering a range of research issues (see, for example, Pattichis, 1999; Tang, 2001).

The third difference between this study and that of Gani (1998) is that we use improved proxies for political instability. Gani (1998: 65) speculates that a possible reason for his finding that political instability was statistically insignificant is that his coup dummy variable might not be capturing the intended effect. A problem with Gani's (1998) analysis is that in constructing a coup dummy, he assigns the value of one for years prior to the first coup in 1987 and zero for the rest of the period. This is econometrically incorrect. The shock should be the other way around with the coup dummy taking the value zero for the years prior to 1987 and one thereafter. We use two proxies for political instability. First, we employ a properly constructed coup dummy capturing the correct shock. Second, we use a democracy index for Fiji compiled by the Freedom House, which measures the level of political freedom.

The fourth difference between this study and that of Gani (1998) is that we explicitly test the stability of the long-run parameters. This is the first time this has been done in the literature on the determinants of immigration flows. Hansen (1992) warns that if the parameters of the model are unstable, the results will be unreliable. Consequently, we apply Hansen's (1992) tests for parameter stability. We corroborate the results with a recent test for parameter stability advocated by Pesaran and Pesaran (1997).

The balance of the article is set out as follows. The next section discusses the migration problem in Fiji and gives some descriptive statistics. It is then followed by the empirical specification, the econometric methodology, and the results. Foreshadowing the main results, we find that in the long run all variables are statistically insignificant. In the short run, in contrast to Gani's (1998) findings, political instability is consistently the most important determinant of migration flows while standard of living and real wage differentials are statistically insignificant across all specifications. The penultimate section reports the results of tests for parameter stability and the final section summarizes the findings.

## MAGNITUDE AND COMPOSITION OF EMIGRATION FROM FIJI

Table 1 provides an overview of trends in outward migration from Fiji. The coups in 1987 were a turning point in the level of emigration. Between 1978 and 1986, annual average emigration was 2,300; however, from 1987 to 2000 annual average migration increased to 5,000. Table 1 shows that Indo-Fijians make up most of the emigrants leaving Fiji. Between 1978 and 1986 Indo-Fijians accounted for 83.8 per cent of total Fiji emigration and this figure increased to 88.9 per cent between 1987 and 2000.

TABLE 1

TRENDS IN MIGRATION

| Period | Total emigration from Fiji | | Indo-Fijian emigration | | | Fiji-New Zealand migration | |
|---|---|---|---|---|---|---|---|
| | Number | Annual average | Number | Per cent of total Fiji emigration | Annual average | Number | Per cent of total Fiji emigration |
| 1978-1986 | 20,703 | 2,300 | 17,358 | 83.8 | 1,929 | 1,586 | 7.7 |
| 1987-1996 | 50,050 | 5,005 | 44,643 | 89.2 | 4,464 | 10,566 | 21.1 |
| 1997 | 4,493 | _ | 3,999 | 89.0 | _ | 972 | 21.6 |
| 1998 | 4,829 | _ | 4,273 | 88.5 | _ | 1,237 | 25.6 |
| 1999 | 4,837 | _ | 4,244 | 87.7 | _ | 1,323 | 27.4 |
| 2000 | 5,275 | _ | 4591 | 87.1 | _ | 1,689 | 32.0 |
| 1997-2000 | 19,434 | 4,858 | 14,837 | 88.2 | 4,239 | 5,221 | 26.9 |
| 1987-2000 | 69,484 | 4,953 | 59,480 | 88.9 | 4,406 | 15,787 | 22.7 |

Source: Figures for total emigration and Indo-Fijian emigration are from Mohanty (2001: 57). The 2000 figure is from Kumar and Prasad (2002: 11). Figures for Fiji-New Zealand migration are from the Fiji Bureau of Statistics.

Mohanty (2001: 58-59) identifies four waves of Indo-Fijian outward migration. The first wave was following independence (1970-1977). In this period the principal push factor was uncertainty surrounding the implications of getting independence. The second wave was in 1986 and 1987, which was due to changes in immigration laws in New Zealand that made it easier for Indo-Fijian professionals to migrate. The third and fourth waves of Indo-Fijian emigration followed the 1987 and 2000 coups and were due primarily to political uncertainty coupled with mounting insecurity over agricultural land leases in the late 1990s.

Several commentators have argued that the main factor driving the third and fourth waves of Indo-Fijian migration is that the 1987 coups fractured the racial goodwill underpinning Fiji's multi-racial democracy and altered the political and social make-up of Fijian society (Naidu, 1988; Bedford, 1989). The coups resulted in weaker law and order, increased damage to property, and racially motivated violence, including hooliganism, mainly targeted at the Indo-Fijian population (see Naidu, 1988). Consequently, Indo-Fijians see the emigration issue in Fiji as politically conceived, where emigration is seen as having strong cultural and political disadvantages for one ethnic group, while containing advantages for another (Mohanty, 2001: 56-57).

The Fijian Government has used the fact that a high proportion of those immigrating are ethnic Indian to justify an affirmative action programme where it allocates increased funding to indigenous Fijians. The Fijian Government has tried offering scholarships to ethnic Fijians, reserved half the places in educational institutions for ethnic Fijians, and set up a commission to encourage indigenous Fijians to undertake further education (*Economic Times of India*, 2002). The argument mounted by the Fijian Government is that it is better to invest in indigenous Fijians because the bulk will remain in Fiji after completing education and training and not migrate (Agence France Presse, 2001). The high proportion of Indo-Fijians emigrating is also altering the racial composition of Fiji's population. In 1987 Indo-Fijians were the dominant ethnic group, accounting for 48 per cent of the population, while indigenous Fijians accounted for 46 per cent of the population (Gani, 2000). By 2002, Indians constituted just 44 per cent of Fiji's population while indigenous Fijians accounted for 54 per cent. Based on current migration trends, it is estimated that in the next decade, the proportion of Indo-Fijians will fall to 30 per cent (*Economic Times of India*, 2002).

Table 1 also shows absolute numbers of Fijians migrating to New Zealand as well as Fiji-New Zealand migration as a percentage of total Fiji emigration. Traditionally, most of Fiji's emigrants go to Australia, Canada, New Zealand, or the United States, with these four countries taking around 95 per cent. Before the 1987 coups, New Zealand took a relatively small proportion of emigrants from Fiji. From 1978 to 1986 New Zealand accepted 7.7 per cent of Fiji migrants, but changes in New Zealand's immigration laws in 1986 and 1987 coupled with the effects of the 1987 coups have seen New Zealand become a more important destination. Between 1987 and 2000 New Zealand accounted for 22.7 per cent of emigrants from Fiji.

Of the approximately 75,000 emigrants from Fiji between 1978 and 2001, about 11 per cent were classified as professionals, technical, or related workers. This means that Fiji lost some 9,000 professionals, technical, and related workers over this period (Mohanty, 2002a: 363). Table 2 gives a breakdown of the human capital loss between 1987 and 1999. Over this period Fiji lost just fewer than 7,000 professionals and skilled workers. Teachers are the single biggest group of professionals Fiji has been losing, accounting for more than 30 per cent of skilled emigration. Of the other categories, architects and engineers made up 21 per cent of skilled emigration, accountants 16 per cent, and medical workers 13 per cent. In the last category, the fact that Fiji is losing its nurses in large numbers is of particular concern for the standard of medical services in Fiji. One report suggests that in 1999 149 nurses of all ranks migrated, in 2000 the comparable number was 124, and in 2001 the figure was 130 (*Daily Post*, 2003).

TABLE 2

LOSS OF HUMAN CAPITAL FROM FIJI
BY PROFESSIONAL AND TECHNICAL CATEGORY

| Category | Total loss, 1987-1999 | Percentage loss |
|---|---|---|
| Architects* | 1,439 | 20.9 |
| Accountants | 1,065 | 15.5 |
| Teachers | 2,125 | 30.9 |
| Medical workers** | 893 | 13.0 |
| Other | 1,347 | 19.6 |
| Total | 6,896 | 100.0 |

Notes: *Includes architects, engineers, and related technicians; **includes medical, dental, veterinary, and related workers.

Source: Mohanty (2002b: 364).

## EMPIRICAL SPECIFICATION

Following Gani (1998) and most other modern economic analysis of the migration decision, the empirical specification is based on Sjaastad's (1962) human capital model of migration. Sjaastad (1962) argued that a person will migrate if the present value of expected increased earnings exceeds the present value of investment costs. In practice, instead of attempting to directly measure the net present value of migration, empirical researchers have utilized a regression equation where the number of migrants is specified as a function of variables measuring earnings and costs of migration. The model of the determinants of migration takes the following form:

(1) $$\ln M_t = \alpha_0 + \alpha_1 \ln Y_t + \alpha_2 \ln W_t + \alpha_3 \ln U_t + \alpha_4 \ln T_t + \alpha_5 \ln PI_t + \alpha_6 Time_t + \varepsilon_t$$

Two specifications of the dependent variable have been used in the empirical literature. Akbar and Deverotz (1993) use the absolute number of migrants, while Brosnan and Poot (1987) and Gani (1998) use migrants as a proportion of the population in the origin country. Allowing for these different equations we estimated two versions of equation (1). In the first specification *M* is the total number of migrants from Fiji to New Zealand and in the second specification *M* is the total number of migrants from Fiji to New Zealand as a proportion of Fiji's population.

*W* is the real average weekly wage in New Zealand less the real average weekly wage in Fiji (both in US$). The real wage in New Zealand is about three to four times higher than in Fiji and the differential has increased over time. We expect that as the wage differential between New Zealand and Fiji widens, it will have a positive impact on migration from Fiji to New Zealand. *U* is the unemployment rate in Fiji less the unemployment rate in New Zealand. We use unpublished statistics on unemployment in Fiji obtained from the Fiji Bureau of Statistics.[2]

The sign on the unemployment variable a priori is unclear. The unemployment rate in Fiji has been consistently higher than the comparable figures for New Zealand. Therefore, if migrants leave Fiji in search of a job in New Zealand, the level of migration from Fiji to New Zealand will be positively related to the difference in unemployment rates. However, offsetting this, Fiji has a large informal sector. Most people who are unable to find jobs in Fiji will find it difficult to emigrate because they cannot afford the high cost. Thus, when the unemployment rate in Fiji is high, many of these unemployed workers are absorbed into the informal sector (Prasad, 1998). Moreover, skilled professionals are more attractive to destination countries and, as such, professionals with skills in demand will find it easier to immigrate than those who are unemployed. New Zealand's current goal is "to allow entry to migrants who would make the highest contribution to employment and income growth" (Winkelmann, 2001: 8). This focus of immigration policy makes it much easier for Indo-Fijians with skills and money to invest to move to New Zealand (Brake, 1993).

*Y* is the standard of living differential. The standard of living is proxied by real per capita income in New Zealand less real per capita income in Fiji (both in US$). Similar to real wages and unemployment, there is a significant gap between the per capita incomes of New Zealand and Fiji and the income differential has shown a slight increase over time. We expect that as the standard of living in New Zealand improves in relation to Fiji, more people will migrate from Fiji to New Zealand.

*T* is the cost of migration. The costs are potentially broad including the monetary costs of moving, the opportunity cost of income foregone while moving, and the psychic displeasure felt from leaving family and friends (Gani, 1998: 61). Because it is difficult to adequately capture all these costs in a time series analysis, we follow the extant literature and restrict our measure of cost to the real cost of transport. In our analysis transport costs are proxied by the one way economy class airfare from Nadi in Fiji to Auckland in New Zealand. We expect that the real cost of transport will have a negative effect on the decision to migrate from Fiji to New Zealand.

*PI* represents political instability. In contrast to the effects of political instability since the first coups in 1987 on Fiji's economic and social composition, including a reduction in amenities, New Zealand has a long tradition of participatory democracy going back to the 1890s. As indicated in the introduction, we use two proxies for political instability. First, we proxy *PI* with a dummy variable, which takes the value of one for 1987 to 2000 and zero otherwise. Second, we use the democracy index compiled by Freedom House. Freedom House constructs their democracy indexes with the assistance of local and international printed materials, field visits, and other communications with informed observers. Following a check-list of various components of democracy, countries are assigned a value for political rights between one (most free politically) and seven (least free). Thus, an increase in the value of the index represents lower levels of political freedom. The Freedom House index has been widely used in previous studies to proxy political instability for a wide range of countries including Fiji (for a Fiji study see Gounder, 2002). We expect that increases in political instability coupled with lower levels of political freedom will have a positive effect on migration from Fiji to New Zealand. Finally, a time trend is used to capture secular trends in migration and $\varepsilon$ is a normally distributed error term.

Equation 1 is estimated using annual data for the period from 1972 to 2001, because while data on migration from Fiji to New Zealand and on most of the explanatory variables are available for a longer time period, the democracy index is only available for the 1972 to 2001 period. The democracy index is sourced from Gastil et al. (various). The data on migration levels from Fiji, Fiji's population, real average weekly income, and unemployment in Fiji are sourced from the Fiji Bureau of Statistics. Real weekly wages in New Zealand are extracted from the New Zealand Yearbook while the unemployment rate is sourced from the New Zealand Department of Statistics. Real per capita gross domestic incomes for both countries are extracted from the World Bank World Tables. Finally, the one-way economy class airfare from Nadi in Fiji to Auckland in New Zealand is extracted from the *ABC World Airways Guide/OAG World Airways Guide (Red Book)*, published by the Reed Group.

## COINTEGRATION METHODOLOGY

As discussed in the introduction, we use the bounds testing approach to cointegration within an ARDL framework. The augmented ARDL ($p, q_1, q_2, \ldots, q_k$)

model can be written as displayed below (Pesaran and Pesaran, 1997: 397-399, Pesaran et al., 2001):

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

where

(3) $$\Omega(L,p) = 1 - \Omega_1\delta_1 L^1 - \Omega_2\delta_2 L^2 - \ldots - \Omega_p L^p,$$

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

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

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

(5) $$\beta_i = \frac{\hat{\beta}_i(1,\hat{q}_i)}{\Omega(1,\hat{p})} = \frac{\hat{\beta}_{i0} + \hat{\beta}_{i1} + \ldots + \hat{\beta}_{i\hat{q}}}{1 - \hat{\Omega}_1 - \hat{\Omega}_2 - \ldots - \hat{\Omega}_{\hat{p}}}, \; i = 1,2,\ldots,k$$

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

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

where $\hat{\delta}(\hat{p},\hat{q}_1,\hat{q}_2,\ldots,\hat{q}_k)$ denotes the OLS estimate of $\delta$ 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 $y_t$, $x_{1t}$, $x_{2t}$,...,$x_{kt}$ and $w_t$:

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

where $ECM_t$ is the correction term defined by:

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

Here $\Delta$ is the first difference operator; $\Omega_j^*$, $\beta_{ij}^*$ and $\delta'$ 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 in equation (1). This is based on estimating error correction models by treating each variable, in turn, as a dependent variable. For example, when emigration is taken as a dependent variable the Unrestricted Error Correction Mechanism (UECM) is of the form:

(9) $$\Delta\ln M_t = a_{0M} + \sum_{i=1}^{n} b_{iM}\Delta\ln M_{t-i} + \sum_{i=0}^{n} c_{iM}\Delta\ln Y_{t-i} + \sum_{i=0}^{n} d_{iM}\Delta\ln W_{t-i} + \sum_{i=0}^{n} e_{iM}\Delta\ln U_{t-i} + \sum_{i=0}^{n} f_{iM}\Delta\ln T_{t-i} + \sum_{i=0}^{n} g_{iM}\Delta\ln PI_{t-i} + h_{iM}Time_t + \lambda_{1M}\ln M_{t-i} + \lambda_{2M}\ln Y_{t-i} + \lambda_{3M}\ln W_{t-i} + \lambda_{4M}\ln U_{t-i} + \lambda_{5M}\ln T_{t-i} + \lambda_{^M}\ln PI_{t-i} + \varepsilon_{1t}$$

Similarly, by taking the wage differential, unemployment differential, living standard differential, real cost of transport, and democracy index as the dependent variable, error correction models can be constructed. The $F$ test is used for testing the existence of a long-run relationship. When a long-run relationship exists, the $F$ test indicates which variable should be normalized. Taking the variables in Equation (9) as an example, the null hypothesis of no cointegration amongst the variables in Equation (9) is $(H_0 : \lambda_{1M} = \lambda_{2M} = \lambda_{3M} = \lambda_{4M} = \lambda_{5M} = \lambda_{6M} = 0)$ against the alternative hypothesis $(H_1 : \lambda_{1M} \neq \lambda_{2M} \neq \lambda_{3M} \neq \lambda_{4M} \neq \lambda_{5M} \neq \lambda_{5M} \neq 0)$. This can also be denoted as follows: $F_M\,(M\,|Y\,,W\,,U\,,T\,,P\,I\,)$. The $F$ test for the null hypothesis of no cointegration amongst the variables in other Error Correction Mechanisms can be constructed in a similar fashion to that described here for Equation (9).

The $F$ test has a non-standard distribution which depends upon whether variables included in the ARDL model are $I(0)$ or $I(1)$ and the number of regressors. Two sets of critical values are reported in Pesaran and Pesaran (1997), which provide critical value bounds for all classifications of the regressors into purely $I(1)$, purely $I(0)$, or mutually co-integrated. 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 $F_M(.)$ 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 appropriate lag selection criteria such as the Schwartz Bayesian Criteria (SBC) and in the second step the selected model is estimated by OLS.

## EMPIRICAL RESULTS

In the first step of the ARDL analysis we tested for the presence of long-run relationships in Equation (1). As we use annual data, the maximum number of lags in the ARDL was set equal to two. The calculated F-statistics are reported in Table 3. In both versions of Equation (1), $F_M(.)$ is higher than the upper bound critical value. In model 1 where the absolute number of immigrants is the dependent variable, the calculated F-statistic of 5.134 is greater than the upper critical value bound of 4.329 at the 5 per cent significance level. In model 2 where migration as a proportion of the population in Fiji is the dependent variable, the calculated F-statistic of 5.496 is greater than the upper critical value bound of 5.331 at the 1 per cent significance level. Thus, the null hypothesis of no cointegration cannot be accepted and there is a long-run cointegration relationship amongst the variables in each of the models.

Once we established that a long-run cointegration relationship existed, Equation (1) was estimated using the following ARDL $(m, n, p, q, r, s)$ specification:

$$\ln M_t = \alpha_0 + \sum_{i=1}^{m} \alpha_1 \ln M_{t-i} + \sum_{i=0}^{n} \alpha_2 \ln Y_{t-i} + \sum_{i=0}^{p} \alpha_3 \ln W_{t-i} + \sum_{i=0}^{q} \alpha_4 \ln U_{t-i} + \sum_{i=0}^{r} \alpha_5 \ln T_{t-i} + \sum_{i=0}^{s} \alpha_6 \ln PI_{t-i} + \alpha_7 Time + \varepsilon_t \tag{10}$$

In estimating Equation (10) a maximum of 2 lags was used ($i_{max}$=2). The estimated models presented here are based on minimizing the SBC. The empirical results for each of the models, obtained through normalizing on the log of migration and log of migration as a proportion of population, in the long run are reported in Table 4. Equation (1) and hence (10) were also estimated without the $PI$ variable.

The test for cointegration was also conducted excluding the *PI* variable. The findings indicated the existence of a long-run relationship among the variables without the *PI* (democracy index) variable. For brevity, these results are not presented here. The existence of cointegration allows the short-run model to be estimated. In doing so, we use the coup dummy variable as a proxy for *PI*. The empirical results for each of the models in the short run, together with diagnostic tests are presented in Table 5.

TABLE 3

F-STATISTICS FOR CO-INTEGRATION RELATIONSHIP

| | Critical value bounds of the F-statistic | | | | | |
|---|---|---|---|---|---|---|
| | 90% level | | 95% level | | 99% level | |
| *k* | $I(0)$ | $I(1)$ | $I(0)$ | $I(1)$ | $I(0)$ | $I(1)$ |
| 5 | 2.782 | 3.827 | 3.189 | 4.329 | 4.011 | 5.331 |

| | Calculated F-statistics | |
|---|---|---|
| | Model 1 (dependent variable is $\ln M_t$) | Model 2 (dependent variable is $\ln(M_t/P_t)$) |
| $F_M(M\|Y,W,U,T,PI)$ | 5.1347 | 5.4960 |
| $F_Y(Y\|M,W,U,T,PI)$ | 2.7897 | 2.8511 |
| $F_W(W\|Y,M,U,T,PI)$ | 0.9442 | 0.8689 |
| $F_U(U\|Y,W,M,T,PI)$ | 3.0555 | 3.1773 |
| $F_T(T\|Y,W,U,M,PI)$ | 1.6375 | 1.6820 |
| $F_{PI}(PI\|Y,W,U,T,M)$ | 2.6515 | 2.7032 |

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

Beginning with the long-run results in Table 4 all variables are statistically insignificant, although with the expected signs. Using the Engle-Granger (1987) framework, Gani (1998) found that in the long run the living standards differential and real cost of transport was statistically insignificant, while the unemployment and wage differentials were statistically significant. This is not surprising given that the Engle-Granger (1987) method tends to give unreliable results with finite samples, in particular when applied to a model with more than one explanatory variable.

Turning to the short-run results in Table 5, the error term $EC_{t-1}$ is statistically significant for each of the models with a negative sign, confirming that a long run equilibrium relationship exists between the variables. The error correction co-efficient varies between -0.223 and -0.333, which suggests that once shocked convergence to equilibrium in each of the models is moderate. In each case, the short-run models pass all of the diagnostic tests for autocorrelation, functional form, normality of the residuals, and heteroskedasticity. The fit of the models in each case is relatively good, although it is much better with the democracy index than the coup dummy variable. The better fit with the democracy index could reflect that it conveys more information. The availability of seven classes to rank countries gives the survey more flexibility to capture subtleties, which cannot be picked up in a dummy variable that takes the value one or zero. In a time series framework, this flexibility makes it easier to take account of moderate fluctuations in political freedoms on a year to year basis.

TABLE 4

ESTIMATED LONG-RUN COEFFICIENTS USING THE ARDL APPROACH

| Regressor | Model 1 (dependent variable is $\ln M_t$) | Model 2 (dependent variable is $\ln(M_t/P_t)$) |
|---|---|---|
| $\ln Y_t$ | 5.9078 (1.0051) | 6.3995 (0.9874) |
| $\ln W_t$ | 0.1137 (0.0792) | 0.1837 (0.1141) |
| $\ln U_t$ | -3.1541 (-1.4758) | -3.6460 (-1.4779) |
| $\ln T_t$ | -3.4151 (-0.8525) | -3.9093 (-0.8723) |
| $\ln PI_t$ | 0.4263 (0.3121) | 0.2006 (0.1301) |
| Time | 0.2376** (2.3029) | 0.2585** (2.2146) |
| Constant | -27.2529 (-0.4093) | -41.6482 (-0.5798) |

Note: **Denotes statistical significance at the 5 per cent level.

The living standard and wage differentials are consistently statistically insignificant across all short-run models. Gani (1998) also found that the living standard differential was statistically insignificant, while the wage differential was statistically significant in his short-run model. As Gani (1998: 65) points out, it is possible that differences in per capita real income are too narrow to measure differences in living standards and therefore it does not properly reflect the effect of differences in living standards on the migration decision. As Lam (2002) said, another factor is that while the real average wage is higher in the receiving country, migrants' expectations of future earnings often go unrealized. In some cases this is because their qualifications are not recognized in the receiving country or they face other forms of discrimination in the labour market. Supporting this contention, a study by the Auckland City Council (2001) found that immigrants from Asia-Pacific countries faced higher unemployment rates and received lower wages in Auckland compared with New Zealand-born workers with the same experience and qualifications.

While not directly comparable, previous studies, which have used reduced-form supply-demand models to analyse the determinants of skilled migration have used real income in the destination country to capture the effect of the expected improvement in economic welfare on the migration decision. These studies have mixed results for this variable. Akbar and Devoretz (1993) and Gani and Ward (1995) found that real income in Canada (from 1976 to 1986) and New Zealand respectively was a positive and statistically significant determinant of skilled migration. However, Devoretz and Maki (1983) and MacPhee and Hassan (1990) found that real income in Canada (for the period 1966 to 1973) and the United States respectively was not a statistically significant determinant of skilled migration.

The real cost of transport has the expected sign in all short-run models and is statistically significant at the 5 per cent level in both of the models, which use the coup dummy variable to proxy political instability. This latter result differs from Gani (1998) where the cost of migration is statistically insignificant with the wrong sign. Our finding for the models with the coup dummy is consistent with Brosnan and Poot's (1987) study of Trans-Tasman migration, which also found the cost of migration to have a negative and statistically significant effect on migration.

The unemployment differential has a negative sign and is statistically significant in the models where the democracy index proxies political instability. On the face of it our result for the unemployment differential differs from Gani (1998). Gani (1998) also finds the unemployment differential to be negative and statistically significant. However, Gani (1998) purports to measure the unemployment differential as the unemployment rate in New Zealand less the unemployment rate in Fiji. If the unemployment differential is measured in this way, a negative sign on the unemployment differential is consistent with Fijians migrating to New Zealand because there are more opportunities to secure a job in New Zealand. We measure the unemployment differential as the unemployment rate in Fiji less the unemployment rate in New Zealand because the unemployment rate in Fiji has been consistently higher than the unemployment rate in New Zealand.[3] Thus, in measuring the unemployment differential in the manner, Gani (1998) gives negative values. This precludes taking the natural log and thus rules out using a log specification. What is puzzling, though, is that according to Gani's (1998) empirical specification he also takes the natural log of the unemployment differential (Gani 1998: 59). It is not clear how he does this given how he constructs the unemployment differential.

TABLE 5

SHORT-RUN COEFFICIENTS

| Regressors | Model 1 | Model 1 (without democracy index) | Model 2 | Model 2 (without democracy index) |
|---|---|---|---|---|
| $\Delta lnY_t$ | 1.9278<br>(1.0505) | 0.3723<br>(0.2537) | 1.9661<br>(1.0354) | 0.3261<br>(0.2208) |
| $\Delta lnW_t$ | 0.0553<br>(0.1193) | 1.3409<br>(1.2230) | 0.0378<br>(0.0814) | 1.3039<br>(1.1818) |
| $\Delta lnU_t$ | -1.0983**<br>(-2.1429) | -0.3175<br>(-0.9622) | -1.0497*<br>(-1.9948) | -0.2645<br>(-0.7964) |
| $\Delta lnT_t$ | -1.1777<br>(-1.1924) | -1.8221**<br>(-2.3446) | -1.1365<br>(-1.1232) | -1.8348**<br>(-2.3491) |
| $\Delta\Delta PI_t$<br>(Democracy index) | 1.0110***<br>(2.7953) | - | 1.0261***<br>(2.7631) | - |
| $\Delta ln PI_{t-1}$<br>(Democracy index) | 0.8025**<br>(2.6251) | - | 0.7543**<br>(2.3918) | - |
| $Coup_t$ | - | 0.4388**<br>(2.4419) | - | 0.4255**<br>(2.3503) |
| $ECM_{t-1}$ | -0.3012*<br>(-1.7438) | -0.2234**<br>(-2.1451) | -0.3328*<br>(-1.8222) | -0.2341**<br>(-2.2078) |
| Constant | -12.5462<br>(-0.5621) | -0.2847<br>(-1.5440) | -9.0697<br>(-0.4022) | -0.2668<br>(-1.4370) |
| Diagnostic tests | | | | |
| $R^2$ | 0.9056 | 0.4222 | 0.9001 | 0.4150 |
| $\bar{R}^2$ | 0.8584 | 0.2777 | 0.8501 | 0.2688 |
| $\chi^2_{Auto}(1)$ | 0.8328 | 1.0102 | 0.8290 | 0.7403 |
| $\chi^2_{Norm}(2)$ | 0.0961 | 1.3464 | 0.2073 | 1.2206 |
| $\chi^2_{Hetero}(1)$ | 0.1896 | 0.0565 | 0.5322 | 0.0093 |
| $\chi^2_{RESET}(1)$ | 0.6934 | 0.0097 | 0.1122 | 0.0034 |

Notes: *indicates statistical significance at 10% level; **indicates statistical significance at the 5% level; ***indicates statistical significance at the 1% level. The critical values for $\chi^2(1)$=3.84 and $\chi^2(2)$=5.99 at the 5 % significance level.

In contrast to Gani (1998), both proxies for political instability are statistically significant with the expected sign. The democracy index is significant at 1 per cent in both the current period and with a one-period lag. The result for the one-period lag is consistent with the fact that migration is often associated with a time lag. The finding for political instability is consistent with Gani and Ward (1995) who find that political instability is significant and positively related to the migration flow of professionals from Fiji to New Zealand. It is also consistent with the findings of Huang (1987) and Lam (2002) for other countries. Huang (1987) found that an index of political and civil rights was the major factor explaining indirect skilled migration to the United States, while the effect of income differentials was quite small. Lam (2002) found that lack of political confidence in the Chinese Government was the major determinant of outward migration from Hong Kong following the end of British rule in 1997.

## CONSTANCY OF COINTEGRATION SPACE

Hansen (1992) cautions that estimated parameters of a time series may vary over time. Thus, parameter tests are important, since unstable parameters can result in the model being misspecified, which has the potential to bias the results. To test for parameter stability we use the Pesaran and Pesaran (1997) and Hansen (1992) tests.

According to Pesaran and Pesaran (1997), the short-run dynamics are essential in testing for the stability of the long-run coefficients. The Pesaran and Pesaran (1997) test amounts to estimating the following error correction model:

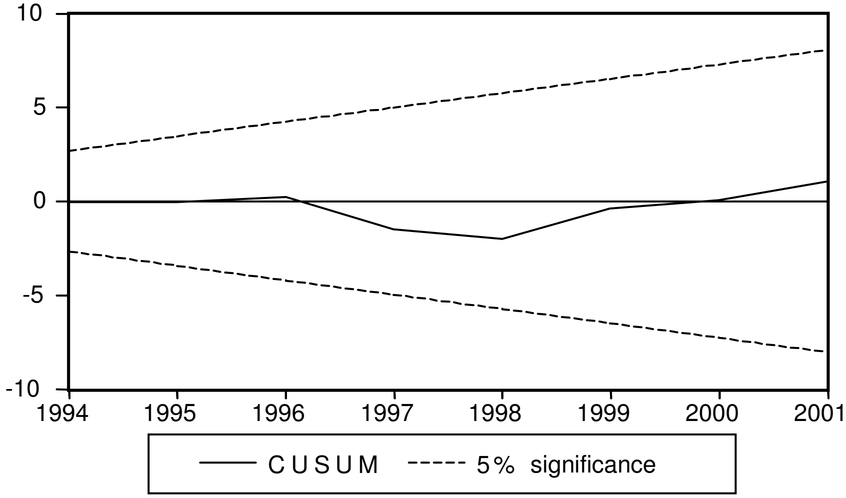
$$\Delta \ln M_t = \alpha_0 + \sum_{i=1}^{m} \alpha_1 \Delta \ln M_{t-i} + \sum_{i=0}^{n} \alpha_2 \Delta \ln Y_{t-i} + \sum_{i=0}^{p} \alpha_3 \Delta \ln W_{t-i} + \sum_{i=0}^{q} \alpha_4 \Delta \ln U_{t-i} + \sum_{i=0}^{r} \alpha_5 \Delta \ln T_{t-i} + \sum_{i=0}^{s} \alpha_6 \Delta \ln PI_{t-i} + \kappa ECM_{t-1} + \varepsilon_t \tag{11}$$

Once Equation (11) has been estimated, Pesaran and Pesaran (1997) suggest applying the cumulative sum of recursive residuals (CUSUM) and the CUSUM square (CUSUMSQ) tests proposed by Brown et al. (1975) to assess the parameter constancy. Equation (11) was estimated by OLS and the residuals were subjected to the CUSUM and CUSUMSQ test. Figures 1 and 2 plot the CUSUM and CUSUMSQ statistics for the two versions of equation (11). The results clearly indicate parameter constancy since the plot of the CUSUM and CUSUMSQ statistics are confined within the 5 per cent critical bounds of parameter stability.

As a further check on parameter stability, the parameter non-constancy tests for $I(1)$ processes advocated by Hansen (1992) were employed. Hansen (1992) proposes three tests – *SupF*, *MeanF*, and $L_c$ – which all have the same null hypothesis 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].

FIGURE 1

CUSUM AND CUSUM SQUARES PLOT FOR MODEL 1



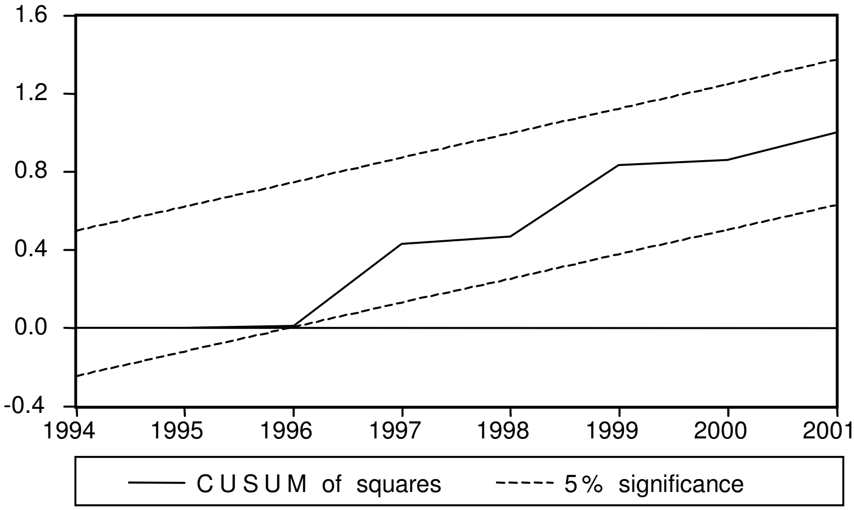
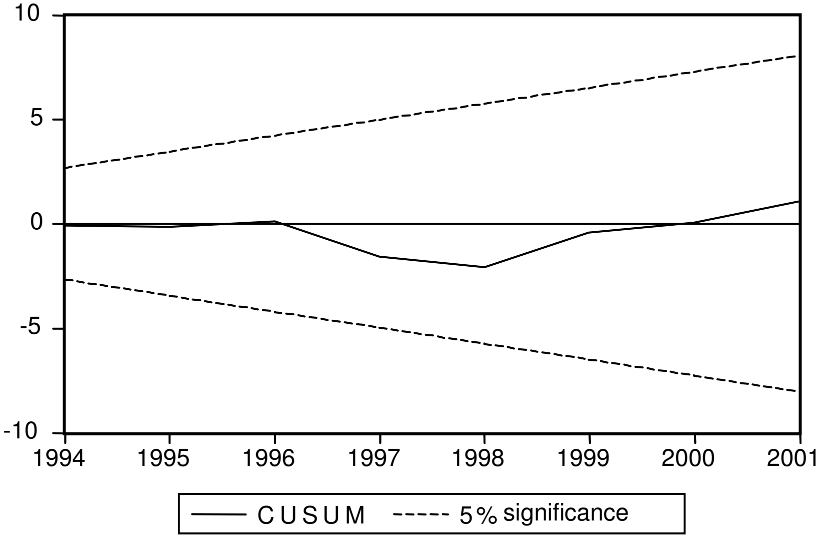
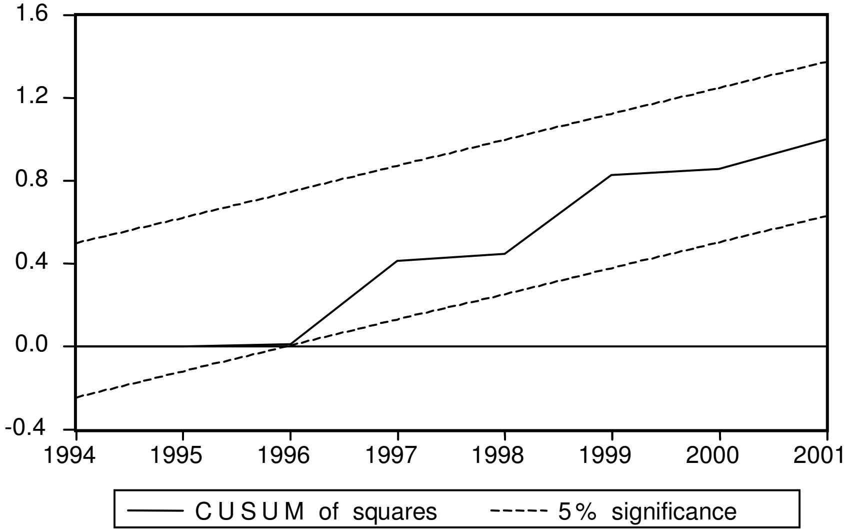



FIGURE 2

CUSUM AND CUSUM SQUARES PLOT FOR MODEL 2



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 6. They indicate parameter stability, since the probability values for each test, for both versions of equation (11), are greater than 0.05. Thus, the results of the Hansen test statistics are consistent with those of the Pesaran and Pesaran (1997) test.

TABLE 6

HANSEN TESTS FOR PARAMETER STABILITY

| Tests | Test statistic | Probability value |
|---|---|---|
| Model 1 | | |
| $L_C$ | 0.0681 | >0.20 |
| *MeanF* | 0.7319 | >0.20 |
| *SupF* | 1.9508 | >0.20 |
| Model 2 | | |
| $L_C$ | 0.0648 | >0.20 |
| *MeanF* | 0.6976 | >0.20 |
| *SupF* | 1.8043 | >0.20 |

Note: The test programme is available at http//www.ssc.wisc.edu/bhansen/.

## CONCLUSIONS AND POLICY IMPLICATIONS

In this article we examine the short-run and long-run determinants of migration from Fiji to New Zealand. The main motivation for the study was to reconsider the findings of an earlier study by Gani (1998) on the same subject. The present study differs from Gani (1998) in several respects. The most important difference is that we use a more recent cointegration technique – the bounds testing approach – that provides more robust results in finite sample sizes than the Engle Granger (1987) approach used by Gani (1998). Moreover, we have suggested that Gani's (1998) construction of the coup dummy variable was incorrect for he assigns the value one for the pre-coup years and zero for the coup years. In addition to correcting this, we also use a democracy index for Fiji to proxy political instability. Further, given that our data set is longer, it enables us to model the impact of the 2000 coup in Fiji. Lastly on the methodological front, we explicitly tested the stability of our model using the Hansen (1992) suite of tests and the Pesaran and Pesaran (1997) test for parameter stability.

Our statistical tests indicate that the variables in our migration model are co-integrated. This is corroborated by the finding that the error correction term is negatively signed and statistically significant. Among our key results, in sharp contrast to the finding of Gani (1998), and as expected, we find that political instability has a positive and statistically significant impact on migration from Fiji to New Zealand in the short run. This has important policy implications for Fiji. The high rates of professional and skilled migration have been a concern for policy makers for more than a decade.

At a conceptual level emigration of skilled labour could have a positive or negative effect on real income in Fiji. One argument is that if highly skilled workers raise the productivity of the economy, their emigration causes a loss. This results in an inward shift in the economy's production possibility frontier, reducing the country's productive capacity and causing a decline in economic growth (RBF, 2002: 41). The human capital loss would also include the cost of spending on the education and training of migrants, the loss of financial capital flows with migrants, and the opportunity costs incurred in educating more people to replace them. An alternative argument is that if expected income in the foreign country is higher, emigration opportunities give people more incentive to acquire skills. Thus, human capital accumulation could end up being higher for those who remain in the home country (Mountford, 1997; Biene et al., 2001; Ueda, 2002). Moreover, the home country could gain through remittances from emigrants generated through large-scale human resource outflow, which in turn contributes to a higher rate of economic growth.

The consensus view is that emigration from Fiji has had an adverse effect on real income in Fiji (House, 2001; RBF, 2002). One report suggests that on average Fiji is losing $F44.5 million per annum through emigration from expenditure on education, training, and immigrants' transfers and legacies alone (Fijilive, 2002). Because Fiji's human capital base was relatively low to begin with, skilled migration running at just over one-tenth of total migration is making a large dent in Fiji's remaining human capital reserves. The International Labour Organization suggests that half of Fiji's stock of middle and high level workers have left Fiji (House, 2001). The phenomenon of "brain drain" is placing a strain on Fiji's ability to provide public services through two avenues (RBF, 2002: 41). First, it reduces the revenue base through decreases in tax collection. This invariably affects, among other things, funding available for education, health, and law and order. Second, brain drain has undermined the quality of public services for those remaining in Fiji, in particular in education and medicine. This is also creating a high opportunity cost of training people to replace skilled emigrants who have left Fiji.

Emigration is also hurting economic growth through reducing savings available for investment. Between 1987 and 2001 it is estimated that, in addition to personal savings, emigrating Fijians withdrew around $F242 million from the Fiji National Pension Fund (RBF, 2002). Moreover, while some other South Pacific countries, such as Samoa and Tonga, have benefited from remittances which have contributed positively to economic growth, in Fiji's case the level of remittances have been non-existent or negligible (Mohanty, 2002b: 365). Political instability, coupled with high rates of migration, is also having an adverse effect on economic growth through reducing investor confidence. Several previous studies have found that political instability has had a statistically significant negative effect on GDP growth in Fiji (see e.g. Gounder, 1999, 2001, 2002; Chand, 2000; Narayan and Smyth, 2004). The Reserve Bank of Fiji (RBF) has set a target of a 5 per cent real growth rate, which requires private investment between 20 per cent and 25 per cent of real GDP. Yet, between 1987 and 2000, average private investment was a mere 5 per cent.

This is having a negative impact on the standard of living in Fiji. In 1998, Fiji ranked forty-fourth in the world in the UNDP Human Development Index. In 2000, it had fallen to sixty-sixth out of 174 countries (RBF, 2002). In light of this, it is essential that Fiji retain its skilled and professional labour. This is particularly so with migration to New Zealand expected to increase with ongoing political instability, following a recent announcement by the New Zealand Government that it intends to introduce an additional migration quota for Fiji citizens over and above current intake levels under its Pacific Access Scheme. This is expected to make it easier for Fijians to migrate to New Zealand (*Daily Post*, 2002). Our empirical results suggest that if Fiji is to stem the tide of outward migration to New Zealand and other developed countries, political stability must rank high as a priority. Political conflicts should be resolved within the confines of the Constitution, which would curtail fears of insecurity and contribute to confidence in the nation. We believe that confidence grows with equal distribution of resources, equal opportunities for employment, and non-discriminatory policies, something that has unfortunately not been present in Fiji since the first military coups in 1987.

## NOTES

1. Helpful comments and suggestions from Ganeshwar Chand, Seema Narayan, and two anonymous referees of this journal on earlier versions of this article are acknowledged. However, any errors or omissions are our own doing.
2. These figures are higher than the official unemployment statistics published by the Fiji Bureau of Statistics. The unpublished statistics are much closer to estimates of the actual unemployment rate by academics and others who have long argued that Fiji's official unemployment statistics understate the true amount. Chand (1998), for instance, estimated unemployment to be 20 per cent in 1998.

3. It is worth noting that Gani (1998) uses the official unemployment rate in Fiji, which is lower than the unpublished estimates of Fiji's unemployment that we use. The official unemployment rate in Fiji, however, is still consistently higher than the unemployment rate in New Zealand.

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## LES FACTEURS DETERMINANTS DE L'IMMIGRATION DES ILES FIDJI VERS LA NOUVELLE-ZELANDE: UNE REEVALUATION EMPIRIQUE SELON LA METHODE DE L'APPROXIMATION DES LIMITES

Cet article réexamine les conclusions de Gani (1998) sur les facteurs déterminants des flux migratoires entre les îles Fidji et la Nouvelle-Zélande en appliquant à la co-intégration la méthode dite du "bounds testing" dans un processus auto-régressif. Il en résulte pour l'essentiel que, sur le long terme, toutes les variables sont statistiquement insignifiantes, quoique correctement signées, exception faite du différentiel du chômage. A court terme, contrairement aux conclusions de Gani (1998), la stabilité politique apparaît de manière insistante comme le facteur déterminant le plus important des flux migratoires, tandis que le niveau de vie et l'écart entre les niveaux de revenus réels sont statistiquement insignifiants dans toutes les spécifications.

## LOS FACTORES DETERMINANTES DE LA INMIGRACIÓN DESDE LAS ISLAS FIJI HACIA NUEVA ZELANDIA: UNA REVALUACIÓN EMPÍRICA QUE RECURRE A LAS PRUEBAS DE APROXIMACIÓN DE LOS LÍMITES

Este artículo reexamina las conclusiones de Gani (1998) sobre los factores determinantes de las corrientes de migrantes desde las Islas Fiji hacia Nueva Zelandia al utilizar el procedimiento de aproximación de los límites para la cointegración, dentro de un modelo autoregresivo de rezagos distribuidos. Los principales resultados apuntan a que a largo plazo todas las variables son insignificantes desde el punto de vista de las estadísticas, aunque bien señaladas con la excepción de la diferencia sobre el desempleo. A corto plazo, existe un gran contraste con los resultados de Gani (1998), la inestabilidad política es constantemente el factor más determinante en las corrientes migratorias, mientras que el nivel de vida y las verdaderas diferencias salariales son insignificantes desde el punto de vista de estadísticas en todas estas especificaciones.