This paper examines the long-run and short-run determinants of migration from Fiji to Australia between 1972 and 2001 using a human capital framework, which is extended to take account of political instability in Fiji. Our main findings are that in the long run the real wage differential and political instability in Fiji are the main determinants. In the short run, there is some evidence that the wage differential and transport costs are important factors, but this finding is not robust across all specifications. Lagged migration and political instability are the most important determinants in the short run.

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Modelling immigration flows: an application of the bounds test to Fiji–Australia migration
Paresh Narayan and Russell Smyth

There is a growing interest in the study of immigration from Fiji among both demographers and economists (see, for example, Gani and Ward 1995; Gani 1998; 2000; Mohanty 2001, 2002). Fiji is an interesting case because in addition to the economic and social factors that influence the migration decision, it experienced two coups in 1987 and a further coup in 2000 that have led to less political freedom and increased political instability. The raw figures suggest that the coups have resulted in a substantial increase in the number of migrants seeking residence elsewhere.

The official statistics show that between 1987 and 2001 75,800 Fiji citizens migrated, an annual average of over 5,000 people (Reserve Bank of Fiji 2002:40). This is almost 10 per cent of Fiji’s entire population, which was just 775,000 in 2001. There was a further sharp increase in immigration following the May 2000 coup. Between May 2000 and October 2002, 16,200 Fijians immigrated, including 14,169 Indo-Fijians (Fiji Labour Party 2003). A further 1,800 Fijians immigrated in the first three months of 2003 (‘Emigration continues to rise’, Daily Post, 11 June 2003). Of even more concern than the
absolute numbers is the fact that many of these migrants are skilled professionals (Naidu 1997; Mohanty 2002). The World Bank (1995), among others, has expressed increasing concern about the effect of the scale of migration of skilled labour on Fiji’s human capital base.

In the 1970s and early 1980s the majority of Fiji immigrants went to North America. However, since the 1987 coups, Australasia has become more important and now Australia is the second major destination for immigrants from Fiji, following the United States. In 1997, about 95 per cent of immigrants from Fiji went to four countries. The United States accounted for 32.7 per cent, Australia for 31.3 per cent, New Zealand for 20.3 per cent and Canada for 10.3 per cent (Mohanty 2001). Reflecting this change, the number of Fijians living in Australia has increased. The 1986 Census recorded that there were 14,756 Fiji-born people living in Australia. The number had increased to 30,544 by the 1991 Census (Stanwix and Connell 1995). Estimates suggest that the total Fiji-born population living in Australia is now in excess of 50,000 people, most of whom are Indo-Fijian (Mohanty 2001).

The increasing importance of Australia as a destination for Fiji immigrants reflects the close economic ties between the two countries, their geographical proximity and Australia’s long history of parliamentary democracy. However, given the deleterious effects which immigration is having on Fiji’s supply of skilled professionals further research is needed into the relative importance of economic and political factors in influencing the decision of Fijians to migrate and also why Australia is becoming such an important destination. There are some demographic studies of migration from Fiji to Australia and Fiji immigrants in Australia, focusing primarily on Fijians in Sydney (see, for example, Munro 1977; Raj 1991; McCall and Connell 1993; Stanwix and Connell 1995). There are, however, no econometric studies of the reasons for migration from Fiji to Australia.

The objective of this paper is to fill this gap through examining the determinants of migrant flows from Fiji to Australia for the period from 1972 to 2000 using a human capital framework, which also takes into account political push factors.

**Related literature**

This article contributes to a growing literature which examines the determinants of migrant flows from low-income to high-income countries. This includes studies of the determinants of migrant flows from low-income countries to the United States (Huang 1987; MacPhee and Hassan 1990); low income countries to Canada (DeVoretz and Maki 1983; Akbar and DeVoretz 1993), from Fiji to New Zealand (Gani and Ward 1995; Gani 1998; Narayan and Smyth 2003) and from the Philippines to the Middle East (Carlos 2002). DeVoretz and Maki (1983) and Akbar and DeVoretz (1993) used reduced-form equations to estimate the determinants of highly skilled immigration to Canada over the period 1968 to 1973 and 1976 to 1986 respectively. DeVoretz and Maki (1983) found that the immigration of professionals from developing countries to Canada from 1968 to 1973 depended on the number of job vacancies in Canada by occupation, competing opportunities in the United States, the total flow of immigrants and the number of people migrating from the immigrant’s home country. In DeVoretz and Maki’s (1983) study the income variable was never significant. However, in the study for the later period, 1976 to 1986, by Akbar and DeVoretz (1993), occupational income was a statistically significant determinant of migration. Akbar and DeVoretz (1993) found that skilled migration from developing
countries to Canada from 1976 to 1986 depended on the supply of new graduates in Canada, occupational income and the previous period’s immigration. Gani and Ward (1995) used the same reduced-form equation approach as DeVoretz and Maki (1983) and Akbar and DeVoretz (1993) to examine the determinants of migration of highly skilled and professionally trained workers from Fiji to New Zealand over the period 1987 to 1990. The explanatory variables which Gani and Ward (1995) included were the previous period’s migration, occupational income in New Zealand, the supply of new graduates in New Zealand and political instability in Fiji. Gani and Ward (1995) found that the number of professional migrants from Fiji to New Zealand was positively and significantly related to the previous period’s migration, real income in New Zealand and political instability in Fiji.

Gani (1998) and Narayan and Smyth (2003) examined the long-run and short-run determinants of migration from Fiji to New Zealand within a human capital framework using time series data and a cointegration and error-correction methodology. Gani (1998) used the Engle and Granger (1987) approach to cointegration and employed time series data for 1970 to 1994. Narayan and Smyth (2003) used the bounds testing approach to cointegration and employed time series data for 1972 to 2001. The empirical specification in both studies posits that the flow of migrants from Fiji to New Zealand is a function of the real wage differential, unemployment differential, standard of living differential, the cost of moving and political instability in Fiji. Gani (1998) found that the real wage and unemployment differentials were the statistically significant determinants of long-run migration from Fiji to New Zealand, while the standard of living differential, the cost of moving and political instability in Fiji were not statistically significant. In contrast to Gani (1998), Narayan and Smyth (2003) found that political instability in Fiji was a statistically significant determinant of migration to New Zealand, consistent with the results from Gani and Ward’s (1995) earlier study for skilled migration.

Other studies have examined the effect of political instability on the migration decision. Lam (2002) used a human capital framework to examine the interaction of economic and political factors on the decision of Hong Kong residents to migrate after Hong Kong was handed back to China in 1997. Lam (2002) hypothesised that migrants leaving Hong Kong would be forced to accept lower incomes in the countries to which they migrated. Lam found that the lack of political confidence in the Chinese government during the transition period had a large positive and significant effect on migration. Lack of economic confidence in the new Chinese regime also had a positive and significant effect on migration, but the magnitude was less than the political effect, whereas the expected decrease in income following migration only had a small negative effect on migration. In a study of professional indirect migration to the United States, Huang (1987) constructed an index of lack of civil and political rights in the migrant’s home country and found that this variable was an important factor in explaining non-returns, whereas the effect of the income differential between the home country and the United States was relatively small.

**Empirical specification**

Our empirical specification is based on the human capital model of migration developed by Sjaastad (1962) and employed by Gani (1998) and Narayan and Smyth (2003). Sjaastad’s (1962) human capital model suggests that an individual will migrate if
the present value of expected increased earnings exceeds the present value of investment costs. Almost all modern economic analyses of migration decisions follow this framework (Lam 2002:489). However, instead of attempting to measure the net present value of migration, existing studies have specified a regression equation where the number of migrants is treated as a function of proxies for the benefits and costs of migration. The empirical specification is as follows

\[ \ln M_t = \alpha_0 + \alpha_1 \ln Y_t + \alpha_2 \ln W_t + \alpha_3 \ln U_t + \alpha_4 \ln T_t + \alpha_5 \ln PI_t + \varepsilon_t \]  

(1)

**Dependent variable**

To measure the dependent variable, the existing literature has used either the absolute number of migrants (Akbar and Deverotz 1993) or migrants as a proportion of the population in the country of origin (Brosnan and Poot 1987; Gani 1998). To accommodate both approaches, we estimate two versions of Equation 1. In the first model \(M\) is defined as the total number of migrants from Fiji to Australia and in the second model \(M\) is defined as the total number of migrants from Fiji to Australia as a proportion of Fiji’s population. The data on migration levels from Fiji to Australia and Fiji’s population are from the Fiji Bureau of Statistics.

**Explanatory variables**

\(Y\) depicts the difference in the standard of living between Australia and Fiji. To proxy the standard of living differential we followed existing studies and used real per capita income in Australia less real per capita income in Fiji (in US dollars). The data on real per-capita income for both countries is from the World Bank’s World Tables. We expect that the number of migrants from Fiji to Australia will be positively correlated with the standard of living differential between Australia and Fiji.

\(W\) represents the wage differential between Australia and Fiji in US dollars. It is measured by the real average weekly wage in Australia less the real average weekly wage in Fiji (in US dollars). The data on real average weekly income in Australia is from the International Monetary Fund’s International Financial Statistics and the data on real average weekly income in Fiji is from the Fiji Bureau of Statistics. The real wage in Australia is approximately four times higher than in Fiji and the wage differential has increased over time. We expect that as the wage differential between Australia and Fiji widens, it will have a positive impact on migration from Fiji to Australia.

\(U\) depicts the unemployment differential between Fiji and Australia, measured as the unemployment rate in Fiji less the unemployment rate in Australia. The data on the unemployment rate in Australia are from the Australian Bureau of Statistics’ Yearbook of Australia. Also used are unpublished statistics on unemployment in Fiji obtained from the Fiji Bureau of Statistics. The unemployment rate in Fiji has been consistently higher than the comparable figures for Australia. We expect that the level of migration from Fiji to Australia will be positively related to the difference in unemployment rates.

\(T\) is designed to capture the costs of migration. The costs of migration can potentially include not only the monetary costs of moving, but also the opportunity cost of income foregone while resettling and the emotional loss of leaving family and friends (Gani 1998:61). However, it is not possible to capture all of these costs in a time-series analysis. Therefore, the existing literature has used the real cost of transport as a proxy for the cost of migration and this article follows this practice. Our proxy for transport costs is the one-way economy class airfare from Nadi, Fiji to Sydney, Australia. The source is the ABC World Airways Guide/OAG World
Airways Guide (Red Book), published by the Reed Group. The real cost of transport is expected to have a negative effect on the decision to migrate from Fiji to Australia.

PI represents political instability. We use two alternative proxies for political instability. First, in the long-run and short-run models we use the Freedom House democracy index (Freedom House, various issues). Freedom House constructs its democracy indexes with the assistance of local and international printed materials, field visits and other communications with informed observers. Following a checklist of various components of democracy, countries are assigned a value of 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. Second, in an alternative specification, in the short-run we proxy PI with a dummy variable, which takes the value of one for the two coup years (1987 and 2000) and zero otherwise. We expect that increases in political instability coupled with lower levels of political freedom will have a positive effect on migration from Fiji to Australia.

Methodology

To examine the long-run relationship between migration from Fiji to Australia and its determinants, we employ the Autoregressive Distributive Lag (ARDL) procedure, also known as the bounds-testing approach to cointegration, developed by Pesaran and Shin (1998) and Pesaran et al., (2001). This procedure can be applied to models irrespective of whether the variables are integrated of order zero (I(0)) or integrated of order one (I(1)). Thus, in contrast to other popular cointegration techniques such as the Johansen and Juselius (1990) and Engle and Granger (1987) approaches, it does not require pre-testing the variables to determine their order of integration. Pre-testing is particularly problematic in the unit-root cointegration literature where the power of unit root tests are typically low, and there is a switch in the distribution function of the test statistics as one or more roots of the $x_t$ process approach unity (Pesaran and Pesaran 1997:184). The ARDL approach, because it draws upon the Unrestricted Error Correction Model, is also likely to have better statistical properties than the Johansen-Juselius and the two-step Engle-Granger method. This is because, unlike the Engle-Granger method the Unrestricted Error Correction Model does not push the short-run dynamics into the residual terms (Banerjee et al. 1993).

The ARDL procedure involves two stages. The first stage is to establish that a long-run relationship exists among the variables in Equation 1. The second stage involves estimating the long-run and short-run coefficients of Equation 1 once it is established that the variables are cointegrated. The mathematical derivation of the long-run and short-run parameters can be found in Pesaran et al. (2001). To implement the bounds test consider a vector of two variables $z_t$ where $z_t = (y_t, x_t)$, $y_t$ is the dependent variable and $x_t$ is a vector of regressors. The data generating process of $z_t$ is a $p$-order vector autoregression. For cointegration analysis it is essential that $\Delta y_t$ be modeled as a conditional error-correction model.

$$\Delta y_t = \beta_0 + \pi_y y_{t-1} + \pi_{yx} x_{t-1} + \sum_{i=1}^{p} \phi_i \Delta y_{t-i} + \sum_{j=0}^{q} \theta_j \Delta x_{t-j} + \theta w_t + \mu_t$$

Here, $\pi_y$ and $\pi_{yx}$ are long-run multipliers, $\beta_0$ is the drift and $\mu_t$ is a vector of exogenous components. Lagged values of $\Delta y_t$ and current and lagged values of $\Delta x_t$ are used to model the short-run dynamic structure. The ARDL test for the absence of any level
relationship between $y_t$ and $x_t$ is through exclusion of the lagged levels variables $y_{t-1}$ and $x_{t-1}$ in Equation 2. Testing for the absence of a conditional level relationship between $y_t$ and $x_t$ entails the following null and alternative hypotheses

$$H_0: \pi_{yy} = 0, \pi_{yx} = 0'$$

(3)

$$H_1: \pi_{yy} \neq 0, \pi_{yx} \neq 0' \text{ or } \pi_{yy} \neq 0' \text{ or } \pi_{yx} = 0'$$

(4)

These hypotheses can be examined using an $F$-statistic. While the distribution of the $F$-statistic is non-standard, Pesaran et al. (2001) report two sets of critical values, which are based on 40,000 replications of a stochastic simulation. This provides critical value bounds for all classifications of the regressors into purely $I(1)$, purely $I(0)$ or mutually cointegrated for a sample size of 1,000 observations (see Pesaran and Pesaran 1997; Pesaran and Shin 1998; and Pesaran et al. 2001 for technical details regarding the computation of critical values for the significance level). We calculated exact critical value bounds for $T=30$ with five regressors, based on 40,000 replications for the $F$-statistic using the Pesaran et al. (2001) GAUSS code.

We included an intercept, but no trend, which is case II in the terminology of Pesaran et al. (2001). The ‘exact’ critical value bounds were as follows: $l(0) 2.407, l(1) 3.517$ (at the 10 per cent level); $l(0) 2.910, l(1) 4.913$ (at the 5 per cent level); $l(0) 4.134, l(1) 5.761$ (at the 1 per cent level). If the computed $F$-statistics fall outside the critical bounds, a conclusive decision can be made regarding cointegration without knowing the order of integration of the regressors. If the estimated $F$-statistic is higher than the upper bound of the critical values then the null hypothesis of no cointegration is rejected. If the estimated $F$-statistic is less than the lower bound of the critical values then the null hypothesis of no cointegration cannot be rejected.

**Empirical results and diagnostic testing**

**Cointegration**

In the first stage of the ARDL analysis we tested for the presence of long-run relationships in Equation 1 treating, in turn, each variable as the dependent variable. For the purposes of this exercise we used the democracy index to proxy political instability. In both versions of Equation 1, $F_M(.)$ is higher than the upper bound critical value at least at the 5 per cent level. In model 1, where the absolute number of immigrants is the dependent variable, the calculated $F$-statistic is 4.981. In model 2 where migration as a proportion of the population in Fiji is the dependent variable the calculated $F$-statistic is 5.814. When the other variables were treated as the dependent variable the $F$-statistic was less than the lower bound critical value at the 1 per cent significance level. Thus, we can conclude that the null hypothesis of no cointegration cannot be accepted and there is a single cointegrating vector. In other words, there is a long-run cointegration relationship amongst the variables in each of the models when the absolute number of migrants (in model 1) or migrants as a proportion of Fiji’s population (in model 2) is treated as the dependent variable.

As we also wanted to estimate the short-run model using the coup dummies rather than the democracy variable to proxy political instability, we additionally tested for cointegration among the variables by excluding the democracy index. The findings indicated the existence of a long-run relationship among the variables without the democracy index. For brevity the results are not reported here, but are available from the authors on request. This result allows the short-run model to be estimated using the coup dummy variables rather than the democracy index as a proxy for $PI$. 

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Long-run results from the ARDL model

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:

\[
\ln M_t = \alpha_0 + \sum_{i=1}^m \alpha_i \ln M_{t-i} + \sum_{i=0}^s \alpha_i \ln Y_{t-i} \\
+ \sum_{i=0}^n \alpha_i \ln W_{t-i} + \sum_{i=0}^q \alpha_i \ln U_{t-i} \\
+ \sum_{i=0}^r \alpha_i \ln T_{t-i} + \sum_{i=0}^s \alpha_i \ln PI_{t-i} + \varepsilon_t
\]

(5)

In estimating Equation 5 a maximum of two lags was used. The estimated models presented here are based on minimising the Schwartz Bayesian Criteria. The empirical results for each of the models, obtained through normalising on the log of migration and log of migration as a proportion of population, in the long run are reported in Table 1. The income and unemployment differentials and real cost of transport are statistically insignificant in both models. However, the real wage differential and political instability, proxied by the democracy index, are statistically significant at the 1 per cent level with the expected signs in both models. A positive sign on the democracy index suggests that less political freedom, reflected in a shift up the index from the democratic end (Freedom House’s 1) towards the authoritarian end (Freedom House’s 7), would lead to more immigration from Fiji to Australia.

To test for stability of the long-run parameters we used the Pesaran and Pesaran (1997) test. Pesaran and Pesaran (1997) posit that the short-run dynamics are essential for testing the stability of the long-run parameters.

### Table 1 Estimated long-run coefficients using the ARDL approach

<table>
<thead>
<tr>
<th>Regressor</th>
<th>Model 1 (Dependent variable is (\ln M))</th>
<th>Model 2 (Dependent variable is (\ln(M/P)))</th>
</tr>
</thead>
<tbody>
<tr>
<td>(\ln Y_t)</td>
<td>-3.4845*</td>
<td>-3.5557*</td>
</tr>
<tr>
<td></td>
<td>(5.1700)</td>
<td>(5.4293)</td>
</tr>
<tr>
<td>(\ln W_t)</td>
<td>3.3380*</td>
<td>3.1342*</td>
</tr>
<tr>
<td></td>
<td>(0.6674)</td>
<td>(0.6467)</td>
</tr>
<tr>
<td>(\ln U_t)</td>
<td>-0.0625*</td>
<td>-0.0533*</td>
</tr>
<tr>
<td></td>
<td>(0.2151)</td>
<td>(0.2039)</td>
</tr>
<tr>
<td>(\ln T_t)</td>
<td>-0.0057*</td>
<td>-0.0629*</td>
</tr>
<tr>
<td></td>
<td>(0.4481)</td>
<td>(0.4257)</td>
</tr>
<tr>
<td>(\ln PI_t)</td>
<td>1.2940*</td>
<td>1.2627*</td>
</tr>
<tr>
<td>(democracy index)</td>
<td>(0.2012)</td>
<td>(0.1919)</td>
</tr>
<tr>
<td>Constant</td>
<td>18.0919*</td>
<td>6.9567**</td>
</tr>
<tr>
<td></td>
<td>(3.2850)</td>
<td>(3.1195)</td>
</tr>
</tbody>
</table>

*Note:* \(*(**\)) denotes statistical significance at the 1 per cent and 5 per cent levels respectively. Standard errors are in parentheses.
coefficients. The Pesaran and Pesaran (1997) test involves estimating the following error-correction model

\[
\Delta \ln M_i = \alpha_0 + \sum_{t=0}^k \alpha_t \Delta \ln M_{i,t-1} + \sum_{t=0}^k \alpha_{t2} \Delta \ln F_{i,t-1} \\
+ \sum_{j=0}^k \alpha_{t3} \Delta \ln W_{j,t-1} + \sum_{j=0}^k \alpha_{t4} \Delta \ln U_{j,t-1} + \sum_{j=0}^k \alpha_{t5} \Delta \ln T_{j,t-1} \\
+ \sum_{j=0}^k \alpha_{t6} \Delta \ln PL_{j,t-1} + \sum_{j=0}^k \alpha_{t7} \Delta \ln ECM_{j,t-1} + \epsilon_{i,t} \tag{6}
\]

Once Equation 6 is estimated using ordinary least squares, 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 parameter constancy. We do not find any evidence of fluctuations of the CUSUM and CUSUMSQ statistics outside the 5 per cent critical bounds of parameter stability, which indicates that the parameters are stable over time.

**Short-run results from the ARDL model**

The results for each of the models in the short run, together with diagnostic tests, are presented in Table 2. Two versions of each model are presented, using the democracy index and coup dummy variables to proxy political instability. The error-correction term is statistically significant at 1 per cent with a negative sign in each case, which confirms that a long-run equilibrium relationship exists among the variables. The diagnostic tests suggest that autocorrelation, functional form, normality of the residuals and heteroskedasticity pose no problem for interpreting the results in any of the specifications and the fit of the models in each case is relatively good.

The lagged dependent variable is statistically significant with a positive coefficient in all specifications. This is consistent with prior expectations. As Greenwood notes, ‘the more persons who have migrated from state i to state j in the past...the greater will be the quantity of information...(and) the current flow of migrants’ (1970:375). This result suggests that previous migrants from Fiji to Australia have a significant role in attracting current migrants via establishing personal networks through family and friend linkages and migrant agency activities. There are several other studies which have found that migration lagged one period has a positive and significant effect on current migration (Devoretz and Mki 1983; MacPhee and Hussan 1990; Akbar and Devoretz 1993; Gani and Ward 1995; Gani 1998).

The wage differential has the expected positive sign and is statistically significant in the specifications with the democracy index as the proxy for political instability, but is statistically insignificant in the specifications with the coup dummy variables as the proxy for political instability. Transport costs and transport costs lagged one period have the expected negative sign across all specifications and are statistically significant in the regressions using the democracy index to proxy political instability. The unemployment differential is statistically insignificant in all regressions. These results can be compared with those of Gani (1998) who examined the determinants of migration from Fiji to New Zealand. Gani found that the unemployment and wage differential were significant determinants of migration from Fiji to New Zealand, while transport costs was a statistically insignificant determinant.

The living standard differential has an unexpected negative sign in each specification and is statistically significant in the specifications with the coup dummies as the proxy for political instability. Previous studies have got mixed results for this variable. Gani (1998) found that the living standards differential was a statistically insignificant determinant of migration from
## Table 2  Short-run coefficients

<table>
<thead>
<tr>
<th>Regressors</th>
<th>Model 1</th>
<th>Model 1 (without the democracy index)</th>
<th>Model 2</th>
<th>Model 2 (without the democracy index)</th>
</tr>
</thead>
<tbody>
<tr>
<td>$\Delta\ln M_{t-1}$</td>
<td>0.6776*</td>
<td>0.4259**</td>
<td>0.6772*</td>
<td>0.4259**</td>
</tr>
<tr>
<td></td>
<td>(0.1296)</td>
<td>(0.1925)</td>
<td>(0.13929)</td>
<td>(0.1681)</td>
</tr>
<tr>
<td>$\Delta\ln Y_t$</td>
<td>-3.0858</td>
<td>-2.9604*</td>
<td>-3.1553</td>
<td>-3.0112*</td>
</tr>
<tr>
<td></td>
<td>(4.2529)</td>
<td>(1.0363)</td>
<td>(4.4072)</td>
<td>(1.0263)</td>
</tr>
<tr>
<td>$\Delta\ln W_t$</td>
<td>3.8150*</td>
<td>5.2083</td>
<td>3.6763*</td>
<td>4.7797</td>
</tr>
<tr>
<td></td>
<td>(1.5220)</td>
<td>(3.4148)</td>
<td>(1.0023)</td>
<td>(3.3833)</td>
</tr>
<tr>
<td>$\Delta\ln U_t$</td>
<td>-0.4501</td>
<td>0.0302</td>
<td>-0.0993</td>
<td>0.0266</td>
</tr>
<tr>
<td></td>
<td>(0.3127)</td>
<td>(0.1969)</td>
<td>(0.1084)</td>
<td>(0.1523)</td>
</tr>
<tr>
<td>$\Delta\ln T_t$</td>
<td>-0.8915*</td>
<td>-0.3517</td>
<td>-0.8954*</td>
<td>-0.3914</td>
</tr>
<tr>
<td></td>
<td>(0.3354)</td>
<td>(0.4478)</td>
<td>(0.3300)</td>
<td>(0.4449)</td>
</tr>
<tr>
<td>$\Delta\ln T_{t-1}$</td>
<td>-0.5893**</td>
<td>-0.2482</td>
<td>-0.5217***</td>
<td>-0.2529</td>
</tr>
<tr>
<td></td>
<td>(0.2967)</td>
<td>(0.4095)</td>
<td>(0.2889)</td>
<td>(0.4060)</td>
</tr>
<tr>
<td>$\Delta\ln Pit$ (democracy index)</td>
<td>0.7945*</td>
<td>-</td>
<td>0.8096*</td>
<td>-</td>
</tr>
<tr>
<td></td>
<td>(0.1744)</td>
<td></td>
<td>(0.1725)</td>
<td></td>
</tr>
<tr>
<td>$\ln PI_t$ (coup dummy variable)</td>
<td>-</td>
<td>0.6878**</td>
<td>-</td>
<td>0.6852**</td>
</tr>
<tr>
<td></td>
<td></td>
<td>(0.2948)</td>
<td></td>
<td>(0.2919)</td>
</tr>
<tr>
<td>ECM$_{t-1}$</td>
<td>-1.4706*</td>
<td>-1.0563*</td>
<td>-1.6684*</td>
<td>-1.0936*</td>
</tr>
<tr>
<td></td>
<td>(0.2005)</td>
<td>(0.2478)</td>
<td>(0.2427)</td>
<td>(0.2469)</td>
</tr>
<tr>
<td>Constant</td>
<td>0.1113</td>
<td>-0.0677</td>
<td>11.6064**</td>
<td>-0.0443</td>
</tr>
<tr>
<td></td>
<td>(0.1436)</td>
<td>(0.1926)</td>
<td>(5.1811)</td>
<td>(0.1910)</td>
</tr>
</tbody>
</table>

### Diagnostic tests

<table>
<thead>
<tr>
<th>Test</th>
<th>Model 1</th>
<th>Model 1 (without the democracy index)</th>
<th>Model 2</th>
<th>Model 2 (without the democracy index)</th>
</tr>
</thead>
<tbody>
<tr>
<td>$R^2$</td>
<td>0.8273</td>
<td>0.6262</td>
<td>0.8325</td>
<td>0.6377</td>
</tr>
<tr>
<td>$\bar{R}^2$</td>
<td>0.7358</td>
<td>0.4601</td>
<td>0.7429</td>
<td>0.4767</td>
</tr>
<tr>
<td>$\chi^2_{AIC}$ (1)</td>
<td>0.9391</td>
<td>0.6133</td>
<td>3.4289</td>
<td>0.6133</td>
</tr>
<tr>
<td>$\chi^2_{Sargan}$ (2)</td>
<td>0.6429</td>
<td>0.9461</td>
<td>1.4016</td>
<td>0.9461</td>
</tr>
<tr>
<td>$\chi^2_{Hetero}$ (1)</td>
<td>0.5394</td>
<td>0.1319</td>
<td>0.1319</td>
<td>0.8821</td>
</tr>
<tr>
<td>$\chi^2_{RESET}$ (1)</td>
<td>0.0587</td>
<td>3.5469</td>
<td>0.8021</td>
<td>3.5469</td>
</tr>
</tbody>
</table>

**Note:** *(***)*** denotes statistical significance at the 1 per cent, 5 per cent and 10 per cent levels respectively. Standard errors are in parenthesis. The critical values for $\chi^2(1)=3.84$ and $\chi^2(2)=5.99$ at the 5 per cent significance level.
Fiji to New Zealand. Among previous studies that have used reduced form equations, Akbar and Devoretz (1993) and Gani and Ward (1995) found that real income in Canada (for the period 1976–86) 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 1966–73) and the United States respectively was not a statistically significant determinant of skilled migration.

There might be two reasons for the result for the living standard differential (see Gani 1998:65). One reason might be that differences in per capita real income are too narrow to measure differences in living standards and therefore do not properly reflect the effect of differences in living standards on the migration decision. Unfortunately, the UNDP Human Development Index for Australia and Fiji, which would be a broader indicator of differences in living standards, is not available in time-series form. A second factor might be that most people who are in a position to migrate already belong to the upper income category of Fiji society and therefore enjoy higher living standards than average Fijians. For these people, at least initially, living standards might decline once they have migrated to Australia as they adjust to living in a new country.

The strongest results are for the political instability variable, with both proxies statistically significant with the expected signs across all specifications. Our results for the democracy variable suggest that a reduction in political freedom, reflected in a move up the Freedom House Index, has had a positive effect on migration. The results for the coup dummies reinforce this finding, suggesting that the coups have had a positive effect on migration. This result is consistent with previous studies that have found that the coups have had a negative effect on Fiji’s economic growth (Gounder 1999, 2001, 2002; Chand 2000). The finding that political instability has increased migration from Fiji to Australia is consistent with most previous studies, which have also found that political uncertainty is a significant factor influencing the decision to migrate. Gani (1998) found that political instability had no effect on migration from Fiji to New Zealand; however, Gani and Ward (1995) and Narayan and Smyth (2003) both found that the coups did increase migration levels. Huang (1987) found that an index of political and civil rights was the major factor explaining indirect skilled migration to the United States. Lam (2002) found that lack of confidence in the Chinese Communist government was the main determinant of outward migration from Hong Kong following the end of British rule in 1997.

Conclusion

This paper has used the ARDL approach to cointegration to model empirically the long-run and short-run determinants of migration from Fiji to Australia. In the long run the real wage differential and political instability in Fiji are the main determinants of migration from Fiji to Australia, while in the short-run lagged migration and political instability are the main factors driving migration. While there is some evidence that transport costs and the wage differential have been important in the short run, the short-run results for the economic variables are not robust across specifications. The clearest results are for political instability with both proxies statistically significant with the expected sign. A growing literature has emerged since the 1987 coups arguing that improved political conditions are required to halt the exodus of immigrants and maintain Fiji’s skilled labour base, which is
needed for sustained economic growth (Naidu 1997; Mohanty 2001, 2002; World Bank 1995). The results presented here provide strong support for the view that heightened political instability has been one of the main reasons for increased Fiji–Australia migration.

We conclude with some brief comments on the policy implications of our findings. Large-scale permanent migration is antithetical to the proper development of the Fiji economy. Fiji’s population represents an important human resource. If Fiji is to retain this important human resource, it should target the factors contributing to migration. In this respect, the policy implications of our long-run results for the wage differential and political instability variables are clear. The findings for the wage differential suggest that one avenue to dissuade potential migrants from out-migration is to increase real wages in Fiji. The findings for the political instability variable suggest that the Fiji government can reduce outward migration through fostering a stable political environment. Naidu (1988) has argued that the 1987 coups ruptured the racial goodwill underpinning Fiji’s multi-racial democracy and changed the social fabric of Fijian society. Given it is the Indo-Fijians who have migrated in large numbers to Australia and because outward immigration in Fiji is seen as having strong cultural and political disadvantages for one ethnic group and advantages for another (Mohanty 2001), it is important that the political environment be made secure for as wide a cross-section of the Fijian population as possible.

Note

1 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, estimates unemployment to be 20 per cent in 1998.

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