IMMIGRATION UNEMPLOYMENT RELATIONSHIP: THE EVIDENCE FROM CANADA*

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This paper examines the relationship between unemployment and immigration in Canada. The bi-directional causality test finds no evidence of a significant effect of Canadian immigration on unemployment. Cointegration tests indicate that there is no observed increase in aggregate unemployment due to immigration in the long run. The results from the causality test based on the vector error correction model confirm that, in the short run, past unemployment does cause (less) immigration but not *vice versa*. There is also a long-run positive relationship among *per-capita* GDP, immigration rate and real wages. The results indicate that, in the short-run, more immigration is possibly associated with attractive Canadian immigration policies, and in the long-run, as the labour market adjusts, Canadian-born workers are likely to benefit from increased migration.

I. Introduction

In recent years there has been renewed public policy and research interest concerning the effects of immigration on countries that both receive and send migrants. Historically, immigration has shaped the economic and population growth rates as well as the demographic composition in many developed countries. However, lower rates of economic growth are associated with high unemployment in many immigrant-receiving countries. At the same time, popular fears about the 'limits to growth' have increased concern about continuing inflows of migrants, particularly from developing and less developed countries.

This is true in Canada as well. At present, the most debatable issue in Canada concerning increased immigration is the impact of foreign worker inflows on the labour market. Some Canadians view that either immigrants steal jobs from Canadian-born workers or that immigrants are less skilled and, owing to difficulties in finding jobs, put pressure on the public purse. The common thread in each of these polar opinions is that immigrants are responsible for unemployment. Others argue that (1) there are skill shortages in Canada and immigrants relieve these bottlenecks, (2) the presumption of a fixed number of jobs in the economy yields a biased conclusion; we need to allow an endogenous determination of both the labour demand and labour supply due to immigration.

This debate between immigration and unemployment in Canada is of particular interest for Australia for several reasons. Australia is a country whose shape has also been profoundly

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influenced by international migration. Australia is one of the major immigrant receiving countries with an immigration population similar to that of Canada. Unlike the USA, where family-sponsored immigration is favoured, both Australian and Canadian immigration policies work on a points basis and favour immigrants with skills. The contention between immigration and unemployment has also been the subject of debate and a matter of public concern in Australia (see Tian & Shan, 1999; Shan *et al*., 1999) as Addison and Worswick (2002) note that 'perhaps as a reflection of this "common view" Australian immigration levels have historically been restricted'.

Immigrants might reduce the job prospects of natives through their adverse effect on the search efficiency of indigenous workers. Depending on their relative ability to find jobs, immigrants may provide strong competition to native-born workers and increase their unemployment. Immigration during recessions can cause unemployment; either directly through the number of new immigrants that are unable to find jobs, or indirectly if migrants displace workers from their existing jobs.

In a general equilibrium framework, immigrants create jobs through the purchase of goods and services independently of their participation in the labour market. This benefits indigenous workers immediately. Immigrants contribute to aggregate expenditure directly through their spending and indirectly through industrial and government expenditure on their behalf, and such expenditures help create employment. This can result in lower unemployment and higher wages for the natives of the country. Immigrants are assumed to have a weak bargaining position in wage negotiations (perhaps stemming from their poor language abilities or the costs they have already paid in order to emigrate), which makes them inpatient to start working. This means that immigrants will accept relatively low wage[1].

The result of this is that firms will now pay a lower average wage and will therefore increase their demand for labour since the cost of hiring a worker declines. Immigrants will be better off as they will now enjoy better employment prospects and higher wages. Native workers can be better off as unemployment might fall through increased job opportunities. Besides, when jobs are plentiful, this ensures that natives have a stronger bargaining position in wage negotiations and can therefore earn higher wages. Hence, immigration possibly makes everyone better off.

The theoretical literature of the effects of immigration on unemployment does not establish any conclusive idea. Harris and Todaro (1970) use a two-sector model of migration and unemployment to describe the possible negative effects of immigration on natives. Ortega (2000) provides a theoretical rationale for positive effects of immigration on natives' wages and unemployment. Berry and Soligo (1969) conclude that the arrival of immigrants generally improves the economic situation of natives. Winter-Embler and Zweimuller (1996) present a two-tier insider-outsider framework where the earnings of natives may be positively linked to the number of immigrants. From the available empirical literature the conclusion is far from clear-cut, for Canada as well as for other developed countries[2]. The net effect of immigration on the labour market largely depends on the particular conditions of these economies and on varying elasticities to immigration. As stressed by Borjas (1994), 'the economic impacts of immigration will vary by time and by place, and can be either beneficial or harmful'.

[1] Firms also anticipate that they will be able to pay lower wages to immigrants because of their high search costs of looking for jobs.

[2] The positive impact of immigration on the labour market for natives for the United States is documented by, for example, Bean *et al*. (1988), Winegarden and Khor (1991), Simon *et al*. (1993), Partridge and Rickman (2006) for Europe by Pischke and Velling (1997) [Germany], Dustmann *et al*. (2005) [Britain], Winter-Embler and Zweimuller (1996) [Austria], for Australia by Withers and Pope (1993), Chapman and Cobb-Clark (1999) and Addison and Worswick (2002). Possible negative impacts of labour market outcome are found by Gross (2002) [France], Angrist and Kugler (2003) [European Union], Lee (1992) [Canada].

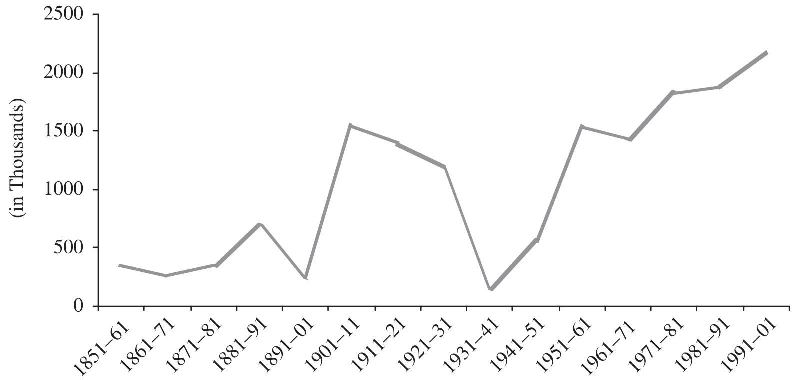

Figure 1. Immigration in Canada: 1851–2001
Source: Immigration Canada <http://www.cic.gc.ca>

In practice, whether a given immigration flow adds or subtracts from the pool of unemployed workers depends on the relative strengths of supply and demand side effects and on the policies that accommodate these flows. Therefore, the economic outcome of immigration is difficult to determine *a priori*, and ultimately becomes an empirical issue. This paper attempts to resolve empirically the relationship between immigration in Canada and its unemployment rate since the 1960s. The objective of this paper is to examine the long- and short-run dynamics of unemployment and immigration.

The paper is organised as follows. The next section describes the trends and patterns of immigration and unemployment in Canada. In Section III, we briefly present the data used in the paper, including selected descriptive statistics. Section IV outlines the empirical strategy. The estimation results are presented in Section V. Concluding remarks are given in Section VI.

II. Immigration and Unemployment in Canada

There are about 100 million people in the world who now reside in a country where they were not born. Canada has experienced a massive flow of immigrants in the last 100 years. In the 1950s and 1960s, immigrant flows originated primarily in Europe, the United States, and the British Isles. After the introduction of the point system in 1967, there has been a strong increase in the proportion of immigrants from Asia; mostly from southern and eastern Asia. In the year 2000, more than 40 per cent of all immigrants in Canada came from only five countries in Asia: China, India, Pakistan, the Philippines, and Korea. In the last two decades, Canada has admitted over 200 000 immigrants per year. This annual flow is not much higher than it was in the late 1950s, the mid-1960s, and the mid-1970s.

Figure 1 illustrates the migrant inflows to Canada since 1851. We see from Figure 1 that there has been a sharp increase in immigration trends since the 1950s. Although the immigrant flow to Canada is significant, the immigration rate (defined as a per cent of the Canadian population) is relatively low in Canada (less than 0.1 per cent). This rate is also low compared to what is observed in the US. In the 1990s immigrant flows to the US were almost 0.3 per cent of the population (Islam, 2005).

Table I Descriptive statistics for key variables

Time period	*Wage (W)*	*SD (W)*	*Immigration rate (M)*	*SD (M)*	*Unemployment rate (U)*	*SD (U)*	*GDP per capita (Y)*	*SD (Y)*
1961–1970	61317.0	310.1	1.76	0.68	4.72	1.06	3894.7	453.6
1971–1980	6901.7	199.9	1.56	0.53	6.44	1.11	5311.6	483.1
1981–1990	7099.2	96.3	1.27	0.41	7.12	1.57	6407.8	498.6
1991–2002	7265.0	292.5	1.90	0.31	5.53	1.14	7367.9	626.3

In the early 1990s, when the Canadian economy was in recession, labour market conditions deteriorated significantly for new immigrants. As a result, immigrants lost considerable ground compared with workers born in Canada. In 1996, only 61 per cent of recent immigrants aged 25 to 44 held jobs, compared with 78.4 per cent of the Canadian born population in the same age group, a gap of 17.4 percentage points. In the first half of the 1980s, a small gap in the employment rate between recent immigrants and Canadian-born workers emerged, and this became extremely pronounced during the early 1990s, reaching a peak in 1996. Census Canada 2001 shows that the gap in labour market conditions between recent immigrants and native born Canadians persisted, despite the strong economic growth of the late 1990s. In 2001, only 65.8 per cent of recent immigrants were employed, 16 percentage points lower than the rate of 81.8 per cent among those born in Canada. The unemployment rate 12.1 per cent for recent immigrants aged 25 to 44 was still twice the rate of 6.4 per cent for the Canadian born population.

III. Data and Descriptive Statistics

The data sources used in this paper include the 2001 Canadian Labour Force Historical Review and DRI Basic Economics Data for (*per capita*) wage (W) and unemployment rates (U) and CANSIM II, from which data for GDP *per capita* (Y) and immigration rates (M) were collected. The immigration rate is expressed as total number of migrant inflows in a given quarter per one thousand Canadians in that quarter. It applies for the movement of nationals who have been granted the right to permanently live in Canada by immigration authorities[3]. The unemployment rate is the percentage of the labour force that actively seeks work but is unable to find work in a particular quarter. Nominal wage data include total compensation paid during the calendar quarter, regardless of when the services were performed, for all employees in Canada. Nominal GDP is an expenditure-based measure and covers the total value of goods and services produced in a given quarter. Both GDP and wages are deflated by GDP deflator. Table I summarises the mean and standard deviation (S.D) of these key variables[4].

[3] It usually applies to persons born outside Canada but may also apply to a small number of persons born inside Canada to parents who are foreign nationals. Note that the migration rate used here differs from 'net migration' rate as the data did not include those individuals departing Canada. However, such a departure rate is very low in Canada and therefore the results are unlikely to change by using the 'net migration rate'.

[4] We conduct the diagnostic checks for the data. The Jarque-Bera (J-B) statistic confirms normality for each series. The J-B statistics for unemployment, the immigration rate, wages and GDP are 10.3, 3.8, 17.2 and 6.3, respectively. Furthermore, there are no Autoregressive Conditional heteroskedasticity (ARCH) effects in the data.

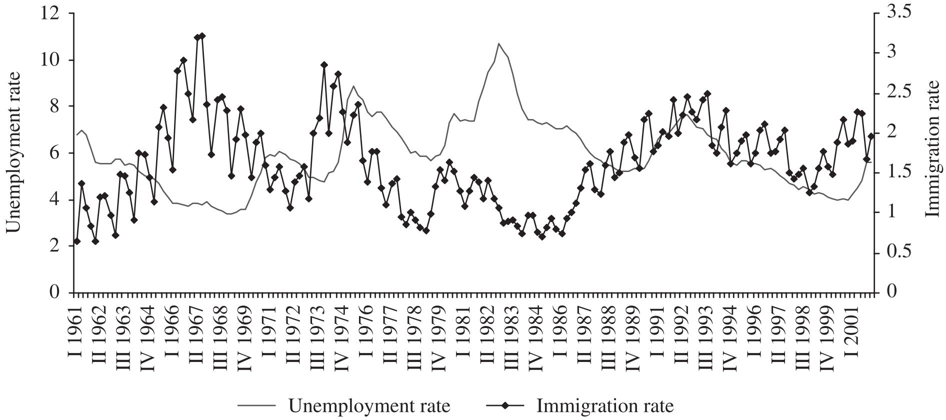

Figure 2. Immigration and unemployment rates in Canada

Figure 2 shows the quarterly unemployment rate and immigration rate in Canada over the period of 1961:1-2002:1. From Figure 2 we observe that the unemployment rate was relatively high in mid-1970s and early 1980s (at the time a major revision in immigration regulations took place) but was more or less declining in the latter half of the 1990s. The unemployment rate again started to rise in the early 1990s but has declined since 1993 (through 2001). In the 1990s, the Canadian immigration policy changed and for the first time in Canadian history, the number of accepted applicants became independent of the state of the economy. However, there was still considerable variation in the immigration rate in the 1990s. From Figure 2 we see that there is apparently no relationship between immigration and unemployment rates. The immigration rate appears to be inconsistent with the observed rise in the unemployment rate in some of the periods.

IV. Empirical Strategy

Economic theory provides no clear basis for predicting the effects of the number of immigrants on unemployment rates among native-born workers. It is generally believed that immigrants find employment in activities where their human capital – with respect to schooling, knowledge of language, and work experience – is more suitable. In a sufficiently segmented labour market, there would be no measurable impact on the indigenous labour force. Alternatively, segmentation might expand job opportunities for native-born workers to the extent that these immigrants constitute a source of complementary inputs. It is at least equally possible, however, to assume that these immigrants share, to an unknown extent, the labour markets occupied by those native-born workers whom they most nearly resemble in terms of human capital. In such a competition, immigrants possibly take jobs away from natives, at least from the marginal workers. In theory, the net outcome is difficult to determine. We therefore examine the causality between immigration and unemployment. In this context, the relevant question is: Is it unemployment that causes immigration or does immigration cause unemployment? In the former case, there is a less politically significant but much debated question in the migration literature as to whether the supply of migrants responds significantly to unemployment, both as a matter of individual migrant choice and as a consequence of government restrictions on inflows.

In the later case, the motivating hypothesis is that migration has effects on both demand and supply, which are difficult to assess *a priori.*

In this context, we want to test the hypothesis as to whether changes in the Canadian immigration rate cause changes in Canada's unemployment or *vice versa*. Withers and Pope (1993) surveyed studies from several countries and time periods. They find that labour market conditions in receiving countries influence intentional migrant inflows, unless other factors intervene. Marr and Siklos (1994) show that before 1978, changes in immigration levels did not affect the Canadian unemployment rate, but after 1978 immigration levels appear to contribute to changes in unemployment. Gross (2004) finds that, in the case of British Columbia, in the short-run, unemployment increases with larger immigration flows, however, in the longer run, unemployment is permanently lowered. It appears that the Granger causality between immigration and unemployment in Canada can go both ways.

This simple Granger causality test implicitly assumes that the information relevant to the prediction of explanatory variables is contained solely in the time series data for these variables. The shortcoming of such causality analysis is that it is merely a statistical exercise with no structural representation in the labour market. There is also an inconclusive debate in the statistical literature as to whether the causality methods can suffer from an omitted variable problem[5]. We therefore specify a more general model that can be used to determine the relationship between immigration and unemployment rates. When the natural rate of unemployment differs across different skill/demographic groups of the labour force, any change in the skill/demographic composition of the labour force will affect the long-run/natural rate of unemployment. Thus, theoretically it is possible that immigration, through changes in the composition of the labour force, will affect the long-run unemployment rate. The expanded model of immigration flows can be written as:

$$y_t = \alpha_0 + \beta_i y_{t-i} + \gamma_i x_{t-i} + \varepsilon_t \tag{1}$$

where y_t is vector consisting of immigration rate, unemployment rate, (*per capita*) wages, to measure aggregate domestic labour market conditions and *per capita* GDP, to measure aggregate domestic economic activity, x is a vector of exogenous variable (seasonal dummies), α_0 is a vector of constant terms, β_i and γ_i are all matrices of parameters of the endogenous and exogenous variables respectively, and $\varepsilon_t \sim$ iid N (0, Σ). Equation (1) is a general equilibrium framework where all the variables of interest are symmetrically and endogenously determined. It captures the supply and demand effects of immigrants as well as feedbacks from wage and the labour supply to determine the final impact of immigrations on the destination market.

The objective here is to develop a strategy for identifying the dynamic adjustment of the market in the short-run and in the long run. So, we focus on to efficiently estimate long-run relationships (if any) between the endogenous variables while identifying the short-run parameters. This is done by adopting the two stage modelling process developed in Johansen (1995), Johansen and Juselius (1994), Granger (1986), and Engle and Granger (1987). In the first stage, we concentrate on cointegration analysis to determine the long-run relationship where the unemployment rate, real wage, real GDP *per capita* and immigration rates are considered to be simultaneously determined. The results can therefore be interpreted as the long-run response of the labour market to immigration. In the second stage, any cointegrating relationship found in the first stage is used in constructing an error correction model (ECM) to identify the short-run dynamic response of unemployment rate, real wages, immigration flows and *per capita* GDP.

[5] An attempt to resolve this issue consists of constructing a theoretical labour market model such as the ones developed by Layard, Nickell and Jackman (1991), Marr and Siklos (1994), and Gross (2002).

Table II Tests for statistical causality between immigration and unemployment

Dependent variable	*Causal variable*	*Causal lag*	*Test statistics*			*Significant individual causal lags at 5% Level*
			F statistic	*Likelihood ratio*	*Wald*	
ln*U*	ln*M*	2	0.491	1.030	0.982	–
ln*U*	ln*M*	4	1.367	5.802	5.467	–
ln*U*	ln*M*	6	1.829	11.756	10.973	2,5
ln*U*	ln*M*	8	1.514	13.303	12.114	2,4,5
ln*U*	ln*M*	10	1.105	12.556	11.053	–
ln*U*	ln*M*	12	0.894	12.592	10.723	–
ln*M*	ln*U*	2	5.201[a]	10.588[a]	10.402[a]	–
ln*M*	ln*U*	4	3.501[a]	14.461[a]	14.003[a]	–
ln*M*	ln*U*	6	2.542[b]	16.113[b]	15.251[b]	–
ln*M*	ln*U*	8	1.874	16.304[b]	14.992	–
ln*M*	ln*U*	10	1.728	19.206[b]	17.281	–
ln*M*	ln*U*	12	1.713	23.288[b]	20.551	3,4

Notes: ln*U*, ln*M* are natural logarithm of unemployment rate and immigration rate respectively. The superscripts, a, b refer to significance at the one and five per cent level respectively.

V. Estimation Results

Estimation of statistical causality

We use seasonally unadjusted data since seasonal adjustments would remove part of the lag sensitivity that is of direct concern in analysing causality. An examination of the data indicates substantial seasonality, both for unemployment and for immigration. Hence, we use quarterly dummy variables to control for seasonality.

Table II outlines the results of the Granger casualty test. The lag length p is selected using multivariate generalisations of the Akaike information criterion (AIC). First, a twelve-quarter lag dependent variable is adopted following AIC. We use F, Wald, and likelihood ratio (LR) tests to test the null hypothesis that the causal lags are jointly zero. The Wald and LR statistics are treated as χ^2 statistics. The F statistic from the regression results suggests the direction of causality from unemployment to immigration at lags 2 and 4, while there does not appear to be any reverse causality from immigration to unemployment. This result is also supported by the Wald test and LR test. There is also causality from unemployment to immigration at higher order lag specifications, though not all of them are supported by the test procedures[6]. The test statistics indicate that the outcome of the Granger causality test is sensitive to the number of lags introduced in the model and to the test used to detect causality[7]. However, the finding that immigration does not cause unemployment also supports the results of studies carried out by Marr and Syklos (1994), Withers and Pope (1985), Tian and Shan (1999) and Shan *et al.* (1999).

We next address the subject of stability. To examine changes in the relationship being considered the standard statistical tests for stability are used. These are the CUSUM and CUSUMQ statistics

[6] Note that although the LR and Wald statistics are both distributed as chi-square variables, they can give different results. Thus, it is possible to reach different conclusions, depending on which is used (see Greene, 2000). However, we find that the Wald test rejects the hypothesis at the ten per cent level of significance for those cases when the LR tests reject at the five per cent. The two statistics, though they have different values, do not differ significantly in terms of absolute value. Unfortunately there is no clear rule of how to proceed when they give us conflicting result.

[7] We can use lag length test, e.g. AIC, LR tests, for Granger causality, but we leave it until we use vector error correction model to test for causality.

Table III Tests for stationarity

Variable	*ADF test statistic*	*Ng-Perron test statistic*
ln*Y*	−2.43	1.07
ln*W*	−2.94	1.23
ln*M*	−1.69	−0.08
ln*U*	−2.78	−2.66

proposed by Brown, Durbin, and Evans (1975) based on the cumulative sum of the recursive residuals. The tests detect parameter instability if the cumulative sum lies outside the area between the two critical lines. These techniques are appropriate in this context due to uncertainty concerning structural changes represented by changes in immigration policy in different time periods, for example, the Vietnam war (1964–73), oil price shocks (1973–74, 1978–79). The stability tests produced no evidence of instability at the conventional level of significance[8].

Cointegration analysis

We use the cointegration technique proposed by Johansen and Juselius (1990, 1994). Our basic model is then the four dimensional vector autoregressive model with Gaussian errors as:

$$y_t = \mu + \sum_{t=1}^{k} \pi_i y_{t-i} + Dx_t + \varepsilon_t, \quad \varepsilon_t \sim iidN(0, \Omega) \quad t = 1, \ldots, T \tag{2}$$

where $y_t = [\ln U, \ln Y, \ln M, \ln W]'$ is a $k = 4$ vector of non-stationary variables, $x_t = [s1, s2, s3]$ is a $d = 3$ vector of centered seasonal dummies, ε_t is a vector of innovations, and μ is vector of constants, π_i is the matrix of coefficients in the ith lag of y_t, and D is the $k \times d$ matrix of coefficients on x_t. We can re-parameterise the model (2) to account for the non-stationarity of the variables in levels. According to Johansen and Juselius (1990, 1994) there exist $k \times r$ matrices α and β each with rank r such that $\pi = \alpha\beta'$ and $\beta' y_t$ can be interpreted as being stationary (even though y_t itself is a non-stationary)[9] among the nonstationary variables. The π matrix conveys information about the long-run relationships between y variables, and the rank of π is the number of linearly independent and stationary linear combinations of the variables studied. Thus, testing for cointegration involves testing for the rank of π matrix r by examining whether the eigenvalues of π are significantly different from zero.

The test for cointegration determines the linearly independent columns of π. Johansen (1988, 1995) and Johansen and Juselius (1990) derive two maximum likelihood statistics for testing the rank of π, and hence the number of cointegrating vectors. In one case the alternative hypothesis is that the rank is k, and the test statistic is known as the *trace statistic*. In the second case, the alternative hypothesis is that the rank is $r + 1$ and the trace statistic is known as the *max statistic*.

Before testing for cointegration, we first determine whether all variables of interest are integrated of order 1, I(1). An examination of the autocorrelation functions indicates that all variables are non-stationary. Formally we use the Augmented Dickey-Fuller (ADF) (Dickey & Fuller, 1979)[10] and Ng-Perron (2001) test statistics. Table III summarises the ADF and

[8] A Chow breakpoint test centered on these periods also supports the stability hypothesis.
[9] In this case, model (2) can be interpreted as an error correction model. See Engle and Granger (1987) and Johansen and Juselius (1990).
[10] An examination of our data indicates that the ln*Y* and ln*W* series contain a linear time trend; ln*M* and ln*U* do not contain a linear time trend. The test results are based on the trend for the former two series, and no trend for the latter two series.

Table IV The eigenvalues and eigenvectors

	Eigenvalues ($\hat{\lambda}$)		
0.1744	0.1136	0.1052	0.0391
	Eigenvectors ($\hat{\Lambda}$)		
1	1	1	1
0.170	0.753	0.068	–5.196
15.756	–3.190	–7.870	–18.020
–23.764	8.228	1.051	34.376

Ng-Perron (2001) test statistics. We choose a four-quarter lag for our quarterly data to determine the stationarity of the natural logarithm of the variables. The test statistics indicate that all the variables have a unit root. After first differencing, the null hypothesis of non-stationarity for each series is rejected at the one or five per cent levels. The ADF and Ng-Perron tests on the second-differences indicate over-differencing, since the coefficient of the lagged level (i.e. the second difference of the series) significantly exceeds minus one in absolute value. So, the test statistics suggest that the stationary series is in first differences.

Next, we consider whether there exists a cointegrating vector[11] between the variables in vector y. For testing the order of the VAR, we use lag length test p using the multivariate generalisations of the AIC. We also use the LR statistic as recommended by Sims (1980). In selecting the number of lag to be included in the model we also need to consider two opposing issues. On one hand, we know that too short a lag length might produce serially correlated errors. On the other hand, a highly over-parameterised model could induce insignificant and inefficient parameters. As indicated by the AIC and LR tests, we opted for a system with a lag length of six, i.e. a VAR (6) model (see Table AI)[12].

On the basis of a plot of the series, a model that allows a linear trend in the variables and in the cointegration space is fitted to the data[13]. The results of the eigenvalue and the eigenvector[14] calculations are given in Table IV.

Let us now consider the number of cointegrating vectors. Table V reports the cointegration analysis for the VAR model with six lags. For the trace statistic, the null hypothesis tests whether the number of cointegrating vectors is less than or equal to r against the general alternative. So, using the λ_{trace} statistic, we reject the null hypothesis of no cointegrating vectors and favour the alternative of one or more cointegrating vectors. Next, we use the $\lambda_{trace(1)}$ statistic to test the null of $r \leq 1$ against the alternative of two or more cointegrating vectors. In this case, the statistic is 43.19, which exceeds the five per cent critical value, but is less than the one per cent critical value, and so a conclusion cannot be reached as to whether the hypothesis should be accepted or rejected. We therefore consider the λ_{max} statistic which allows us to test the null hypothesis of r cointegrating vectors against the specific alternative of $(1 + r)$ cointegrating vectors.

[11] We use a vector of constants and three quarterly dummies to check the seasonality.
[12] The graphs of the residual series of each equation in the VAR also indicate that the residuals are white noise at lag six, thus, fulfilling the standard recommendation to pick the order of the VAR.
[13] See Johansen and Juselius (1990) for a discussion on this issue.
[14] Each column of the β matrix gives an estimate of a cointegrating vector. The cointegrating vector is not identified unless we impose some arbitrary normalisation. So, we adopt the normalisation procedure so that the r cointegrating relations are solved for the first r variables in the vector as a function of the remaining $(k–r)$ variables.

Table V Cointegration test for the system

H_0	H_1	$\lambda_{\max} = -T\ \ln(1-\hat{\lambda}_{r+1})$	*10 % critical value*	*5% critical value*	*1% critical value*
$r = 0$	$r > 0$	73.66	59.14	62.99	70.05
$r \leq 1$	$r > 1$	43.19	39.06	42.44	48.45
$r \leq 2$	$r > 2$	24.02	22.76	25.32	30.45
$r \leq 3$	$r > 3$	6.34	10.49	12.25	16.26
H_0	H_1	$\lambda_{trace} = -T \sum_{i=r+1}^{n} \ln(1-\hat{\lambda}_i)$	*10 % Critical Value*	*5% Critical Value*	*1% Critical Value*
$r = 0$	$r = 1$	30.48	29.12	31.46	36.65
$r = 1$	$r = 2$	19.17	23.11	25.54	30.34
$r = 2$	$r = 3$	17.68	16.85	18.96	23.65
$r = 3$	$r = 4$	6.34	10.49	12.25	16.26

Notes: The asymptotic distribution of the test statistics is given by an expression involving stochastic integrals of Brownian motion (see Johansen, 1988, 1991). The critical values for trace and max statistics are reported from Osterwald-Lenum (1992)

The calculated value of the $\lambda_{max}(0, 1)$ statistic is 30.48, so we can reject the null hypothesis that $r = 0$ against the specific alternative $r = 1$ at the ten per cent level but cannot accept $r = 2$. Thus we conclude that we have only one cointegrating regression. This hypothesis will be maintained below.

Finally, we need to verify that the assumption about the presence of a trend in the variables and in the cointegration space is consistent with the data. We test these by the chi-square statistic as suggested by Johansen (1991, 1995). We find that the calculated value of χ^2 statistic is 10.6. Since the asymptotic distribution of this statistic is $\chi^2(3)$ with three degrees of freedom, it is significant, and the hypothesis that we have linear trend in the variables and cointegration relations is maintained. We also reject the model that restricts the 'constant' to the cointegration space.

Since we have only one cointegrating vector, the maximum likelihood estimator for the single cointegrating relation is given by the first column vector of $\hat{\Lambda}$ in Table IV. Consequently the first column represents the β matrix (vector). Note that estimates in Table IV are the unrestricted estimates because we have analysed the general cointegration model, that is, we did not impose any *priori* assumptions on the cointegrating vectors. With $r = 1$, we can find the cointegrating relation as the first column of the long-run coefficient of β. The estimated long-run equilibrium relationship then can be written as (t-statistic is in parenthesis):

$$\begin{array}{l} log(U_{t-p}) = 75.84 - 0.046t - 0.17log(M_{t-p}) + 15.75log(Y_{t-p}) + 23.76log(W_{t-p}) \\ \qquad\qquad\qquad\quad (1.87) \qquad (-0.607) \qquad\quad (-1.94) \qquad\qquad (1.97) \end{array} \tag{3}$$

We see that there is a (statistically insignificant) negative relationship between unemployment and flows of immigrants in the long run. So, in the long run, Canadian employment is not adversely affected by the admittance of immigrants. While not reported here, there is also a statistically significant positive relationship between *per capita* GDP and the *per capita* wage as well as between *per capita* GDP and immigration rate, real wage and unemployment rate yet a statistically insignificant negative relationship between *per capita* GDP and unemployment rate. These results are robust since they hold consistently across the vectors.

Testing for weak exogeneity

Now we consider whether any of the variables can be considered weakly exogenous when the parameters of interest are long-run parameters. The test for weak exogeneity is the linear

Table VI Weak exogeneity test statistic

Variable	*lnY*	*lnW*	*lnM*	*{lnY, lnW}*
$\chi^2(.)$	6.97	4.52	4.41	9.41
P-value	0.01	0.03	0.04	0.01

restriction on α. If a variable is weakly exogenous, it is possible to condition the short-run model on that variable without any loss of information. The weak exogeneity hypothesis can be tested with a LR test procedure described in Johansen and Juselius (1994). Asymptotically it follows a χ^2 distribution, with the degrees of freedom equal to the number of zero restrictions on the α coefficients. We can write the $\hat{\alpha}$ as:

$$\hat{\alpha}' = (0.001, -0.002, -0.004, -0.002)$$

The low coefficient value of $\hat{\alpha}$ means that slow adjustment towards the equilibrium. Since we include the additional variable ln$Y$ and ln$W$ in our VAR specification, and the corresponding values of α (third and fourth coefficient respectively) are close to zero[15].

So we test whether ln$Y$ and ln$W$ are weakly exogenous for the long-run parameters of the model. First, we test the hypothesis of $\hat{\alpha}_3 = \hat{\alpha}_4 = 0$ and we find no evidence for restricting $\hat{\alpha}_3 = \hat{\alpha}_4 = 0$ (see Table VI). A separate test of whether $\hat{\alpha}_3 = 0$ or $\hat{\alpha}_4 = 0$ also rejects the hypothesis of weak exogeneity. Therefore, we can't exclude ln$Y$, ln$W$, and ln$M$ from the system as neither of these variables can be considered weakly exogenous.

Error-correction models

The evidence of cointegration between variables in our model rules out the possibility of Granger non-causality[16], albeit it does not say anything about the direction of this causal relationship. The application of a vector error correction (VEC) model, in this case, will allow the revelation of the direction of the causality, and also makes it possible to distinguish between short-run and long-run Granger causality. The causality tests are conducted by testing whether all the coefficients of the first difference of each variable are statistically different from zero as a group based on a standard χ^2 test and/or whether the β's of the error-correction model are also significant.

First, a stationarity test confirms that the residuals are I(0). Next we consider each of the multivariate vector error correction models using all lags from one to six. Table VII reports the results from the corresponding Granger causality tests. In the table, the sum of the coefficients is reported in the first line. Table VII illustrates that there is one-way temporal causality whereby past unemployment reduces immigration. However, immigration does not, in general, cause unemployment. This is indicated by a statistically insignificant positive coefficient corresponding to the sum of the lagged migration terms in the unemployment equation.

Furthermore, past immigration has a quantitatively smaller impact on increasing unemployment than past unemployment has on reducing the current immigration rate. We also find that, in the short-run, increases in the wage raise GDP, but an increase in *per capita* GDP is associated with less migration. Thus, although one might think that the increase in *per capita* GDP

[15] If it is zero, disequilibrium in the cointegrating relationship does not feed back onto that variable. However, all the variables are in terms of natural logarithms and a low adjustment coefficient is plausible, as any change in variable should be interpreted in terms of growth of the variable rather than level.
[16] Granger (1986) shows that when two time series are cointegrated there exist a causal relationship in at least one direction.

Table VII Temporal causal test from error correction models

	Coefficient of				
Equation	*EC*	$\Sigma dlnM$	$\Sigma dlnU$	$\Sigma dlnY$	$\Sigma dlnW$
$D\ln M$	−1.022	0.896	−0.166	−0.306	0.449
	(−3.43)[a]	(39.48)[a]	(14.11)[b]	(17.03)[a]	(7.29)
$D\ln U$	−0.148	0.142	0.638	0.933	−3.124
	(−1.44)	(9.78)	(48.86)[a]	(6.83)	(7.83)
$D\ln Y$	0.034	−0.014	−0.026	−0.024	1.234
	(1.06)	(6.94)	(10.58)	(155.67)[a]	(12.10)[c]
$D\ln W$	0.012	−0.033	0.040	0.053	0.505
	(0.59)	(10.94)[c]	(8.43)	(9.21)	(15.30)[a]

Notes: The numbers in parenthesis are chi-squared [J] statistics (where J is the number of restrictions), except the first column of the table, which is t-statistic for the single error-correction term. The superscripts a, b and c indicate significance at the one, five and ten per cent level, respectively.

induces more migration in terms of personal motives, our results indicate that more migration is associated with other reasons possibly due to attractive Canadian immigration policy and/or relatively quick assimilation into Canadian society. In Table VII, EC is an error correction term representing the lag of the single cointegrating residual obtained from the procedure discussed above. It represents the extent of departure from the long-run equilibrium and the coefficient of EC is the short-run adjustment parameter. The signs of the coefficients of the error correction terms for $\ln U$ and $\ln M$ indicate that when the unemployment rate and immigration rates exceed their long-run equilibrium level, they should decrease.

Table VII also suggests that all of the variables are econometrically endogenous; every variable provides explanatory power over and above the other variables. Thus, all of the regression equations possess significant feedback effects between $\ln M$, $\ln U$, $\ln Y$ and $\ln W$[17].

VI. Concluding Remarks

The analysis of causality between unemployment and immigration suggests that migration does not lead to unemployment, although there is evidence that unemployment has a negative effect on migration. The estimated VEC model confirms the results from the bi-variate Granger causality test. The estimation of the multivariate VEC suggests that, in the short-run, there is a one-way causal relationship between unemployment and immigration; that is, higher unemployment causes less immigration in Canada. The long run coefficient estimates suggest that there is no adverse effect on unemployment due to immigration. Thus, even if immigration might contribute to temporary unemployment (as indicated by a positive but statistically insignificant short-run parameter value), the effect disappears in the long run. In fact, in the long run, we can see some positive effects on employment, as indicated by the estimated long-run coefficient matrix. The link from unemployment to immigration in the short-run might not be surprising since over the period prior to the 1980s, and part of 1980s immigration levels tended to fall in recessionary times, though in the more recent periods we have seen an end to that approach to setting goals for the intake of immigrants.

[17] Diagnostic tests on residuals based on tests for normality (J-B tests), ARCH LM test and serial correlation LM test do not give concern about the adequacy of the model.

The long run effect of immigration appears to be that an increase in labour demand is matched by an increase in labour supply caused by immigration. Because the positive coefficient implies that native-born wages rise due to a relative increase in immigration, higher short-term unemployment may result from firms employing fewer Canadian-born workers at these wages. However, in the long run, demand side effect takes place, wages adjust, labour demand is restored and thereby Canadian born workers are benefited. Thus, immigration, as a whole, can be considered a Pareto improvement in the long run with the 'migration economy' Pareto dominating the 'no-migration economy'. Immigrants are (assumed to be) better off by their employment and higher wages relative to their country of origin; natives, as a whole, are better off due to the lower rate of unemployment and the corresponding rise in wages in the long run, possibly owing to the stronger bargaining position of Canadian-born workers in a higher job availability environment.

Composition of immigration flows also matters. Increasing the discrepancy between the skill distribution of immigrants and that of the existing workforce contribute to mitigate some of the adverse effects on unemployment. The adverse consequences of immigration, if any, even in the short-run, can also be mitigated by a screening process focused on average level as well as on the distribution of immigrant skills relative to native workers. The distribution of skills is directly controllable in both Canada, and Australia, through the point system. But it may not be possible to have a perfect match between newly arrived immigrants and native workers and it may not be desirable since labour market conditions change more rapidly than do immigration policies. But the screening process through the point system can certainly play a role by adjusting the point system, even if a small one, in the long-run. Further research in this area could be enhanced through use of aggregate time series data set with information about the skill composition of immigrants. To the author's knowledge, however, such data is not available in Canada for public use.

Overall, the paper does not find evidence in favour of the popular contention of the adverse effects of immigration on the Canadian labour market. As such, one cannot claim that immigrants are responsible for the unemployment rate. Rather immigration has the positive effects of creating employment and higher wages. From a policy point of view the results suggest that Canadian policy-makers should pursue current immigration policies and that decreasing admissions below the current level of immigrants is not desirable. However, it should be noted that the immigration and aggregate unemployment relationship is a difficult one to test. Immigration is an endogenous choice and few opportunities exist where a truly exogenous immigration shock can be observed. Furthermore, unemployment is a macroeconomic phenomenon that needs to be explained by factors that are not necessarily linked to immigration.

Appendix

Table AI Lag length tests for cointegration

lag	*AIC*	*log*
1	−3.93	−6.72
2	−15.97	−12.03
3	−16.71	−12.44
4	−17.10	−12.70
5	−17.44	−12.93
6	−17.87	−13.21
7	−17.89	−13.31
8	−17.77	−13.35

Note: Σ is the determinant of the residual covariance matrix value from unrestricted VAR.

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