The Changing Labour Force Participation of Canadians, 1969-96: Evidence from a Panel of Six Demographic Groups

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After rising almost without interruption for four decades, the aggregate labour force participation rate in Canada fell every year in the first half of the 1990s, slumping to 64.8 per cent in 1995-98 from its 1989 peak of 67.5 per cent. Given the definitional identity linking the employment-population ratio (E) to the unemployment rate (U) and the labour force participation rate (L):

(1) $E = (1 - U)*L$,

the sudden large drop in participation combined with the sustained high level of unemployment account for the depression in Canadian employment in this decade. The percentage of working-age Canadians holding jobs fell from 62.4 per cent in 1989 to 58.2 per cent in 1993; by 1997 it had only recovered to 58.9 per cent.

Today, the central macroeconomic question for Canada is: what proportion of this slump in participation and employment is cyclical in the sense that it can be eliminated without bringing ever-rising inflation? To answer this question, we need first an explanation of labour force participation, and second an explanation of the NAIRU — the critical level of unemployment (U*) below which inflation begins to increase without limit. The participation rate (L) depends on perceived employment opportunities as represented by the employment rate (E), and both it and the NAIRU depend on structural factors (S), such as demographic change and social policy. Formally, we have:

(2) $L = L(E, S)$

(3) $U^* = U^*(S)$.

Once these two functions are well-specified, they can be inserted into equation 1 and combined with population projections to give an estimate of potential, non-cyclical employment and participation. Potential employment can, in turn, be combined with trend labour productivity to yield an estimate of potential output.

In this symposium, the focus is on the aggregate labour force participation shown in equation 2. We set ourselves the task of specifying and estimating this equation directly and deriving some of its implications. To this end, we use a 28-year period of macroeconomic observations (1969-96) for six broad demographic groups: men and women aged 15-24, 25-54, and 55 and over. In contrast with the other two econometric papers in this symposium (Archambault and Grignon, 1999; Beaudry and Lemieux, 1999), our particular angle is to cover all demographic groups, emphasize theoretical consistency as well as empirical flexibility, and pay careful attention to the role of social policy (unemployment insurance, social assistance and the minimum wage) among structural factors.

However, because the aggregate data we use are too coarse and degrees of freedom scarce, we do not attempt to distinguish precisely between the specific impacts on labour force participation of several other structural factors that have been mentioned in the literature, such as rising school attendance, the plateauing of women's participation, better income protection and rising average age for older men and women, and worsening employment prospects among the low-skilled. While we believe our analysis offers a reasonable 'explanation' of past and present, our reliance on trend variables naturally implies that we cannot project our results mechanically into the future. We can 'explore' the future, but not predict it.

Our statistical results attribute the drop in Canadian labour force participation about equally to three factors: very poor macroeconomic conditions, major policy changes in unemployment insurance and the minimum wage, and broad-based structural transformations. In the short run, we find it unlikely that even the best macroeconomic prospects can allow the aggregate participation rate to recover more than one-third of its 2.7-percentage-point drop between 1989 and 1997. Prospects for a recovery of the employment-population ratio are better if one believes — as we do — that equilibrium unemployment has declined since 1989.

The paper is organized as follows. Section 1 develops our theoretical model of labour force participation and presents the data used to test it. Section 2 explains the empirical methodology and reports the estimation results. Section 3 tells how these results explain the recent decline in labour force participation. Section 4 explores their implications for the future. Section 5 summarizes the paper and concludes.

Theory and Measurement

The labour force participation rate of each of the six demographic groups is the percentage (L_i) of group population i who want to work according to Labour Force Survey criteria. Following the standard classical framework, we reason that L_i depends on three sets of factors: individual/family preferences and the demographic composition of the group; economic conditions, including wealth and non-labour income, wage and employment opportunities, attractiveness of schooling, and so on; and income security programs, such as unemployment insurance, social assistance and pension benefits.

Our particular implementation of this general framework will have L_i depend on six influences:

(4) $L_i = L_i(J, W/P, W_m/P, rD/M, B/P, T)$

where:

J = index of job availability,

W/P = average real wage,

W_m/P = average real minimum wage,

rD/M = index of unemployment insurance generosity,

B/P = average real social assistance benefit and

T = quadratic annual time trend.

We provide intuition for the impact of each of these variables on labour force participation.

First, better job opportunities (J) on the demand side raise the subjective probability of successful job search and, hence, the expected payoff to labour force participation. Therefore, L_i must be an increasing function of J. We measure job opportunities by the ratio of Statistics Canada's help-wanted index to the total working-age population. This index is a good instrument because it is highly correlated with the (demand-side) probability of finding a job and appears to be insensitive to (supply-side) participation rate shocks (Archambault and Fortin, 1997).[1] Equations 1 and 2 above make clear alternative indexes of labour market pressure such as the unemployment rate or the employment-population ratio are unacceptable as such, because they are endogenous to the employment-participation behavioural system.[2]

Second, the average real wage (W/P) has an uncertain impact on labour force participation. It is true that, as a payoff to working, a higher real wage should draw more people into the labour force. However, it is also true that a higher real wage allows workers to achieve a given income target by participating in the labour force less frequently or for shorter durations. Moreover, in families where income is shared, the real pay raise obtained by one member can act as a deterrent to labour force participation of other members, by increasing their 'reservation wage.' We measure W/P by dividing the average wage for the entire economy (W) by the consumer price index (P).[3]

Third, the average real minimum wage (W_m/P) also has an uncertain impact on labour force participation. For a given level of the average wage in the economy, a higher minimum wage by itself increases the attractiveness of labour force participation for minimally qualified workers. But it may, at the same time, reduce the perceived probability of successful job search and deter labour force participation by objectively reducing the availability of jobs (Mincer, 1976). We measure W_m/P through division of the weighted average of provincial hourly minimum wages (W_m) by the consumer price index (P).

Fourth, more generous unemployment insurance (UI) regulations and benefits should encourage at least some positive degree of labour force participation. This is because the broad payoff to participation not only includes the wage when working, but also the availability and level of UI compensation when out of work. UI generosity was substantial in Canada in 1972-89. For example, in 1978-89, whenever the regional unemployment rate exceeded 11.5 per cent, the '10/42' rule applied. A minimum qualifying period (M) of 10 weeks triggered eligibility of 42 weeks of maximum benefit duration (D) compensated at the rate (r) of 60 per cent of the previous weekly wage. Thus, potential benefits, represented by rD/M, were 252 per cent of the labour income earned during the minimum qualifying period. This implicit wage subsidy amounted to a powerful incentive to join and stay in the labour force. A measure of the severity of federal restrictions imposed on UI in 1990, 1993, 1994 and 1996 is that, at unchanged levels of regional unemployment rates, the implicit wage subsidy fell to 62 per cent in 1997 from 180 per cent in 1989.

We use the implicit wage subsidy thus defined (rD/M) as our measure of UI generosity (Fortin, 1984; Sargent, 1996). We hypothesize that increased generosity in the 1970s raised labour force participation, and this effect was partly reversed by the restrictions of the 1990s. This wide swing in social policy has been specific to Canada; it has not occurred in the United States. Naturally, a test of this hypothesis requires extending our sample of observations on labour force participation back to the 1960s. We also allow labour force participation to respond gradually over several years to changes in UI generosity.

In calculating the national UI wage subsidy, rD/M, the wage replacement rate, r, is set at the common national value. The minimum work requirement, M, and the maximum benefit duration, D, are calculated national averages not of actual, but of standardized provincial values. These standardized values are those M and D would have taken, given the existing legislation, if provincial unemployment rates had remained constant (and equal to their sample averages) throughout the estimation period. This "instrumentation" procedure makes the standardized values of M and D responsive to legislative changes, but not to cyclical changes in regional unemployment rates. It cuts off the correlation that would otherwise arise between random shocks to labour force participation and the actual UI wage subsidy measure. If not removed, this correlation would bias upwards the estimated impact of UI generosity on participation.

Although greater UI generosity tends to induce more people to enter and stay in the labour force, it may nevertheless also operate as work disincentive for those already participating. Essentially, by reducing the marginal cost of not working, more generous UI discourages participation through negative income and substitution effects.

To summarize, UI has two conflicting effects on labour force participation: it brings more people into the labour force, but it may reduce the extent of participation of those already working. Therefore, the net effect of UI generosity on participation could be positive for some groups and negative for others.[4]

Fifth, an increase in the average real social assistance (SA) benefit should weaken labour force participation by reducing the net payoff to going from welfare to work. However, the estimated impact could be muted by the fact that provinces require employable welfare recipients to be available for work. This could bias their responses to the LFS in favour of participation. We measure B/P through division of the average benefit (B) by the consumer price index (P). The average benefit is in turn obtained through ex post division of the sum of all provincial expenditures on general social assistance by the corresponding total number of recipients.

Sixth, we use an annual time trend (T = 0, 1,..., 27 in 1969, 1970,..., 1996) to capture all other smoothly changing variables affecting labour force participation in the sample period. We make this trend quadratic (and later test whether it is cubic) to allow the growth rate of participation itself to change over time. It will catch social, economic, policy, and other unspecified influences as well. Among the developments that have been mentioned in recent literature are the rising rate of school attendance, the changing roles of men and women in society, the changing demand and supply of skills, the development of private and public pensions, and the rising average age of the 55-plus age group.

Estimation Methodology and Results

We implement the six equations explaining the labour force participation rates of the six demographic groups as follows:

$$(5) \quad \Delta\log(L_i) = \beta_{i0} + \beta_{i1}\,\Delta\log(J) + \beta_{i2}\,\Delta\log(W/P) + \beta_{i3}\,\Delta\log(W_m/P) + \beta_{i4}\,\Delta\log(1{+}r\,D/M) + \beta_{i5}\,\Delta\log(B/P) + \beta_{i6}T + \varepsilon_i.$$

Here, i indexes demographic groups (i = men 15-24, men 25-54, men 55 and over, women 15-24, women 25-54, women 55 and over), ε_i is an additive zero-mean random error, and the β_{ij}'s are the coefficients to be estimated (j = 0, 1,...,6). The annual time subscript is omitted everywhere for simplicity.

The log-linear form of the equations allows the β_{ij} coefficients to be interpreted as elasticities. Our theoretical expectations are that we must have $\beta_{i1} \geq 0$ and $\beta_{i5} \leq 0$. The annual differencing of the equations reflects our attempt to eliminate spurious correlations between jointly trended variables. It is made necessary by the fact all six participation rates display strong upward or downward trends in our 28-year sample.[5] In particular, since the level equations have been assumed to be quadratic in the annual time trend T, the above difference equations will be linear in T. A negative value of β_{i6} means that the rate of change in labour force participation L_i itself falls over time. It is well-known, for example, that the labour force participation rate of middle-aged women increased more and more slowly from the 1970s to the 1990s — so β_{i6} should indeed be negative in this case.

For each difference equation, our sample covers the same 28 years of macroeconomic observations, 1969-96. The basic constraint on this time span comes from social assistance data, which begin in 1968 and end in 1996. All variables except the UI generosity variable are measured in current-year values. Our UI generosity variable is a moving average of current- and past three-year values of the basic index. This distributed-lag specification is motivated by the existing Canadian and international evidence on delayed labour market reaction to changes in unemployment insurance legislation (e.g. Fortin, 1994; the Organization for Economic Co-operation & Development, 1994; Lemieux and MacLeod, 1998). It is empirically supported by minimization of the Akaike information criterion.

Statistical estimation of the six equations is based on Zellner's 'seemingly-unrelated-regression' method. This estimation technique improves upon precision of the traditional 'ordinary-least-squares' method by taking advantage of the likely presence of contemporaneous intercorrelation between the unmeasured ε_i error terms of the equations.[6] More details on variable definitions are provided in the data appendix.

Table 1 reports estimation results for equation 5 that exclude the economy-wide average real wage variable W/P. This exclusion is motivated by prior estimation and tests showing this variable plays a statistically negligible role in every equation.[7] Every other variable has a significant impact on participation for at least two demographic groups.

With R^2 statistics ranging between 0.66 and 0.85, the overall explanatory performance of the estimated equations is good for the younger and middle-aged groups, particularly given that estimation is in annual difference form. However, it is not so good in the case of older groups, where the R^2 statistics fall below 0.30 and much unexplained variation remains. Nevertheless, all equations do well when subjected to a battery of diagnostic tests for serial correlation (Durbin-Watson, Breusch-Pagan), omitted variables (RESET) and structural change (Chow, CUSUM). We are, in particular, unable to detect any significant shift in the regression constant and the time trend coefficient of any equation over the 1990-96 subperiod.[8]

One important implication of statistical stability is that the declining aggregate participation rate of the 1990s should not be attributed to some new structural relationship, but to continuation of the same modes of labour market behaviour (including continuation of the same unspecified trends) as were observed over the previous 25 years. This raises our confidence that the recent participation decline can be understood in terms of the particular outcomes for the explanatory variables in this decade.

Before working through how the equations explain developments in the 1990s, we focus on key aspects of the estimated structure.

Table 1 Annual equations for log changes in labour force participation rates, estimated with panel data for six demographic groups, Canada, 1969-1996

	Demographic Groups					
	15-24		25-54		55 and over	
Explanatory Factors	**Men**	**Women**	**Men**	**Women**	**Men**	**Women**
CONSTANT	0.89*	2.52*	0.06	5.31*	-0.47	1.99*
	(0.44)	(0.53)	(0.12)	(0.34)	(0.73)	(1.10)
INDEX OF JOB AVAILABILITY	0.060*	0.038*	0.014*	0.010*	-0.000	0.017
	(0.008)	(0.010)	(0.002)	(0.006)	(0.014)	(0.020)
AVERAGE REAL MINIMUM WAGE	-0.167*	-0.164*	-0.010	-0.075*	-0.009	0.197*
	(0.040)	(0.048)	(0.011)	(0.031)	(0.066)	(0.099)
AVERAGE UI WAGE SUBSIDY	0.086*	0.086*	-0.002	0.013	-0.004	-0.120*
	(0.022)	(0.027)	(0.006)	(0.017)	(0.037)	(0.056)
AVERAGE REAL SA BENEFIT	-0.015	-0.072	-0.013	-0.111*	-0.117*	-0.144
	(0.040)	(0.048)	(0.011)	(0.031)	(0.066)	(0.099)
TIME TREND	-0.053*	-0.120*	-0.012*	-0.190*	-0.081*	-0.140*
	(0.026)	(0.031)	(0.007)	(0.020)	(0.043)	(0.065)
Summary statistics						
Standard error of regression	0.85	1.03	0.23	0.66	1.41	2.13
R^2 statistic	0.84	0.78	0.66	0.85	0.16	0.26
Durbin-Watson statistic	1.98	2.26	2.62	2.06	2.66	2.94

Note: The theoretical background for the equations and variable definitions are presented in the text and the data appendix. As indicated in text equation 5, all variables except the time trend are expressed in annual log difference form. Regression constants, time trend coefficients and standard errors of regressions are expressed in log points. The other coefficients are interpreted as elasticities. The estimation method is Zellner's iterative seemingly-unrelated-regression method. Numbers in parentheses below the estimated coefficients are their estimated standard errors. An asterisk indicates that the zero-coefficient hypothesis is rejected at the 90 per cent confidence level.

Source: Authors' calculations; see text for explanations.

First, the participation rates of all younger and middle-aged groups respond positively and significantly to cyclical variations in job opportunities (J). But there are large differences in amplitude. The estimated cyclical elasticity of participation ranges from 0.060 for young men to 0.010 for middle-aged women. Basically, younger groups (15-24) respond four times more than middle-aged groups (25-54) of the same gender, and men respond 50 per cent more than women of the same age group. By contrast, the older groups (55 and over) do not appear to change their participation behaviour in response to cyclical changes in job availability.

Second, changes in the minimum wage influence labour force participation among the younger groups and women, but not among middle-aged and older men. This pattern is consistent with the fact almost 90 per cent of minimum-wage workers are young or female. The estimated participation elasticities are negative among the younger groups (-0.16) and middle-aged women (-0.08, half as large); it is positive and quite large among older women (0.20).

Third, the positive response of labour force participation to UI generosity is concentrated among the two younger groups, with significant elasticities of 0.09 for young men and women. The two middle-aged groups and older men do not seem to be affected much by changes in UI regulations. For their part, older women seem to react negatively to UI generosity, with a large and significant elasticity of -0.12.

Fourth, increases in social assistance benefits are found to affect negatively participation both for middle-aged women (among whom the phenomenon of single parenthood is concentrated), and older men. The two estimated elasticities are equal to about -0.1. No significant reaction is detected among the younger groups, middle-aged men and older women. It is important to point out that traditionally most provinces have imposed strict controls on the eligibility of young adults for social assistance and kept young welfare recipients under closer supervision than older ones. Since the extent of controls and supervision is likely to increase with the level of benefits, it is not surprising the labour force participation rates of younger groups turns out to be unresponsive to changing benefit levels.

Fifth, all time trend coefficients have significant negative estimated values. The slowdown has been most pronounced among women of all age groups, who started the period with rapidly increasing participation rates. Based on the estimated coefficients for the constant and the time trend, we estimate the trend participation rate began to decline in the mid-1980s for older women, in the late 1980s for younger women, and in the late 1990s for middle-aged women.[9] The rate also declined throughout the sample period for middle-aged and older men, but much more rapidly among the latter than the former. The participation of younger men followed the same pattern as that of younger women, rising up to the late 1980s and declining afterwards.

The slowing growth rate of trend labour force participation has been a broad-based phenomenon, but the reasons no doubt differ between demographic groups. As explained earlier, our statistical analysis does not attempt to identify the various determinants of those trends across demographic groups, but on the basis of the existing literature the following presumptions look reasonable. First, rising school attendance has depressed participation rates among younger men and women (Archambault and Grignon, 1999). Second, the diffusion of labour force participation as a sociological phenomenon has been approaching maturity among women of all ages (Beaudry and Lemieux, 1999). Third, better income protection and rising average age have played a role among older men and women (Baker and Benjamin, 1997). Fourth, declining employment opportunities and worsening working conditions have led to labour force withdrawals among the low-skilled population in general (Murphy and Topel, 1997).

How do our estimated equations explain the decline in aggregate labour force participation in the 1990s? To answer this question, we use the estimated coefficients reported in Table 1 and simulate the impact of the accumulated change in each explanatory factor on the participation rate of each demographic group in 1990-97. This makes 1989 the base year of the simulation.

We calculate two types of aggregate statistics. First, we sum up the simulated impacts of all explanatory factors on the participation rate of each individual group. The actual change in the group-specific labour force participation rate over the period is by definition the sum of this simulated change and the residual prediction error. Second, we calculate the population-weighted average of the simulated group-specific impacts of each factor to obtain its effect on the participation rate of the entire population. The actual change in the aggregate participation rate over the period (2.7 percentage points) is by definition the sum of these population-weighted averages of factor-specific impacts, prediction errors and the effect of shifting population weights.

Formally, as a weighted average of group participation rates, the aggregate participation rate can be written as:

(6) $L = \Sigma A_i L_i$,

where A_i is the population weight of group i. From this definition, the change in L from 1989 to 1997 can be expressed in exact form as:

(7) $\Delta L = \Sigma \overline{A}_i \Delta L_i + \Sigma \overline{L}_i \Delta A_i$,

provided $\overline{A}_i$ and $\overline{L}_i$ are defined as simple averages of the 1989 and 1997 values of A_i and L_i, respectively.

The first term on the right-hand side of equation 7 is the population-weighted average of the group-specific changes in participation rates,

Table 2 Simulated impacts of cyclical, policy and other structural changes on the labour force participation rates of six demographic groups, accumulated in Canada, 1990-97
(percentage points)

	Demographic Groups						
	15-24		25-54		55 and over		Total[a]
Types of Changes	Men	Women	Men	Women	Men	Women	
JOB AVAILABILITY	-3.4	-2.0	-1.0	-0.6	0.0	-0.2	-1.0
MINIMUM WAGE	-1.5	-1.4	-0.1	-0.7	0.0	0.4	-0.4
UI GENEROSITY	-3.1	-2.8	0.1	-0.5	0.1	1.0	-0.5
SOCIAL ASSISTANCE	0.0	0.2	0.0	0.3	0.2	0.1	0.1
OTHER STRUCTURAL[b]	-2.4	-2.2	-1.7	4.0	-6.8	-1.9	-0.8
TOTAL SIMULATED[c]	-10.4	-8.2	-2.7	2.5	-6.5	-0.6	-2.6
SIMULATION ERROR	0.4	-0.6	0.0	-0.6	1.5	0.0	-0.0
TOTAL ACTUAL[d]	-10.0	-8.8	-2.7	1.9	-5.0	-0.6	-2.7[e]

Notes
a Weighted average of the six columns, with weights equal to group population shares. From left to right, these shares are equal to 0.091, 0.088, 0.285, 0.285, 0.115 and 0.136.
b Simulated impact of time trend. The various unspecified influences subsumed under this catch-all variable include the rising rate of school attendance, the changing roles of men and women in society, the changing demand and supply of skills, the development of pension plans, and the rising average age of the 55-plus age group.
c Sum of five above lines.
d Sum of two above lines.
e Includes the small aggregate impact (-0.1 point) of population shifts among the six demographic groups.

Source: Authors' calculations based on Table 1 data; see text for explanations.

which depend on simulated factor effects and residual prediction errors. The second term is what we have just termed the effect of shifting population weights. (Dugan and Robidoux (1999) call it the demographic composition effect). We calculate this effect generated a small reduction of 0.1 of a percentage point in L in 1990-97.[10]

Table 2 reports the factor- and group-specific simulated impacts on participation rates accumulated over the period. Needless to say, the numbers should be taken as orders of magnitude, not as precise figures. In particular, simulations for older men and women retain the same degree of imprecision as the Table 1 regressions on which they are based.

In interpreting these results, it is important to keep in mind the size and timing of the simulated factor changes. The index of job availability dropped from 100 in 1989 to 40 in 1992, then remained almost unchanged in 1992-97. The average real minimum wage was roughly constant in 1989-91, then increased by 13 per cent in 1991-97. The standardized implicit UI subsidy decreased sharply and continuously from 180 per cent of the wage in 1989 to 62 per cent in 1997. The index of average real SA benefit first increased from 100 in 1989 to 107 in 1992, then fell to 96 in 1996.[11] Two facts stand out: the cyclical drop was concentrated in the first half of the period, policy changes in the second half; and the saw-tooth movement in SA benefits did not result in much net change over the period. The only policy developments that really mattered were the large accumulated changes in UI and the increase in the real minimum wage.

The aggregate factor and group statistics in Table 2 give two unambiguous messages. First, across factors the three types of changes contributed about equally to the decline in the aggregate labour force participation rate. Of the 2.7-percentage-point drop in aggregate participation in 1989-97, cyclical reduction in job availability contributed one point, UI restrictions 0.5 of a point, the higher minimum wage 0.4 of a point, and other structural changes 0.8 of a point. Second, across groups the decline in participation was, on the contrary, uneven. Younger men and women reacted most strongly, with their participation rates falling by 10 and 8.8 points, respecitvely. The rates for middle-aged and older men decreased by 2.7 and five points. The participation rate fell only a bit among older women, and it continued to pull ahead among middle-aged women. However, the increase in the latter case was much smaller in the 1990s (1.9 points) than in the 1980s (14.1 points).[12]

At the detailed level, Table 2 shows the younger groups dominated the decline in participation arising from both cyclical and policy changes. Middle-aged men and, to a lesser extent, middle-aged women also reacted negatively to cyclical changes. Middle-aged and older women showed some sensitivity to policy changes, but in opposite directions. For all groups other than younger men and women, other structural changes were the dominant force behind changes in participation. Trend participation increased rapidly (but, again, at a slowing rate) among middle-aged women, and decreased sharply among older men. Among all other groups, trend participation fell moderately.

Exploring the Future

How informative are our estimated equations on the future outlook of labour force participation? We think they can be useful in projecting cyclical and policy-induced developments, but not very helpful in forecasting other structural changes.

Policy-induced developments are the easiest to handle. We believe the major overhaul of recent years in Canadian social policy has been completed, and the next decade will see a lot more stability in this area. We foresee stable real minimum wages and SA benefits, and no significant additional amendment to the Employment Insurance Act. A further reduction of about 0.1 of a percentage point in the aggregate participation rate will occur as a lagged response to past changes in UI regulations, but no more. This would bring the participation rate decline on account of accumulated UI restrictions to 0.6 of a percentage point. Little else is likely to happen in these three areas. The major policy changes of coming years will instead centre on pension policy, a factor our statistical measurement has included in the 'other structural' category.

Concerning cyclical developments, our estimated equations have established quantitative links between the job availability index and group-specific participation rates. But to form an estimate of how much increase in participation would accompany a full economic recovery requires knowledge of the equilibrium, non-inflationary value of the job availability index. This equilibrium value in 1998 and beyond could be different from the actual 1989 value of the index. To determine its level amounts to specifying the level of the NAIRU, beyond the scope of this paper.

In these circumstances, a practical approach consists of estimating how much the aggregate participation rate and the aggregate employment-population ratio would have to increase from their 1997 levels to achieve a range of realistic levels of the NAIRU. Formally, we simulate equations 1 and 2 in log difference form as follows:

(8) $\Delta\log(E) = \Delta\log(1 - U) + \Delta\log(L)$,

(9) $\Delta\log(L) = \beta_1\Delta\log(E)$ + etc.,

where β_1 is a weighted average of the six estimated β_{i1} coefficients with weights equal to the 1997 labour force shares.[13] These β_{i1} coefficients are obtained from re-estimation of equations 5 with the aggregate employment-population ratio (E) replacing the job availability index (J) as the cyclical variable in every group equation.[14] The 'etc.' in equation 9 stands for the sum of all policy and other structural influences.

We consider equations 8 and 9 as a two-equation system that can be solved out for $\Delta\log(E)$ and $\Delta\log(L)$ as functions of $\Delta\log(1 - U)$ and structural variables. We obtain:

(10) $\Delta\log(E) = [1/(1 - \beta_1)]\Delta\log(1 - U)$ + etc.

(11) $\Delta\log(L) = [\beta_1/(1 - \beta_1)]\Delta\log(1 - U)$ + etc.

Our estimate for β_1 is 0.25. Given the 1997 value of the unemployment rate was 9.2 per cent, equations 10 and 11 can be used to calculate how much higher the employment-population ratio

Table 3 Simulated reactions of the aggregate labour force participation rate and the aggregate employment-population ratio to bringing down the national unemployment rate to various hypothesized NAIRU levels Canada, 1997 basis (percentage points)

Unemployment	Participation Rate		Employment-Population Ratio	
From 9.2 per cent to	Increase From 64.8 per cent to		Increase From 58.9 per cent to	
9 per cent	0.0	64.8	0.2	59.1
8 per cent	0.3	65.1	1.0	59.9
7 per cent	0.5	65.3	1.9	60.9
6 per cent	0.8	65.6	2.8	61.7

Note: The NAIRU is the non-accelerating-inflation rate of unemployment. It is the lowest level of unemployment for which inflation does not tend to rise without limit. The 9.2 per cent, 64.8 per cent and 58.9 per cent figures for the unemployment rate, the participation rate and the employment-population ratio are their 1997 actual values. Their respective 1989 actual values were 7.5 per cent, 67.5 per cent and 62.4 per cent.

Source: Authors' calculations; see text for explanations.

and the participation rate would have been in that year if the unemployment rate had been at its NAIRU level and structural variables had been the same.

Table 3 presents the results of these simulations for four alternative NAIRU values (6 per cent, 7 per cent, 8 per cent and 9 per cent), which cover a wide range of opinion on its true value. The 9 per cent upper limit is a figure preferred by those who think that the Canadian output gap has already been closed and the Canadian economy was already overheating in 1998, a fact that could be temporarily hidden by the current world commodity price deflation.

The 6 per cent lower limit can be rationalized as follows. Starting from the 8 per cent estimate produced by the Bank of Canada Research Department in the late 1980s (Rose, 1988), three structural developments since then can be thought of having reduced equilibrium unemployment. First, the share of the high-unemployment 15-24 age group in the labour force has fallen from 21 per cent to 16 per cent. Given the five-point difference between the youth and adult unemployment rates observed in the high-employment years 1988-89, a mechanical reduction of 0.25 of a percentage point in aggregate unemployment must have followed.

Second, we have just estimated that the string of federal restrictions to UI eligibility and benefits will probably have reduced aggregate labour force participation by 0.6 of a percentage point by the end of the 1990s. If most of this drop has essentially shifted some non-employment from the unemployment category to the out-of-labour-force category, as suggested by the analysis of Card and Riddell (1996), then, based on equation 8, the consequence has been a further reduction of 0.8 of a percentage point in structural unemployment.

Third, the recent performance of the U.S. economy at full employment may indicate that increased domestic and international competitive pressure in labour and product markets arising from deregulation, globalization and freer trade has cut structural unemployment by up to one percentage point in North America. Official estimates of the U.S. NAIRU currently put it in the 5 per cent-5.5 per cent range, which is one percentage point lower than the 6 per cent-6.5 per cent range estimated only 10 years ago (Stiglitz, 1997; United States President, 1998).[15]

Whatever the exact level of the Canadian NAIRU, the key result in Table 3 is that, in any future non-inflationary cyclical recovery, the aggregate participation rate is currently unlikely to rise more than 0.8 of a percentage point. The current equilibrium value of the participation rate would not seem to exceed 65.6 per cent. Similarly, the employment-population ratio is unlikely to rise more than 2.8 percentage points in any non-inflationary cyclical recovery. Its current equilibrium value would not seem to exceed 61.7 per cent.

In conjunction with working-age population numbers, this constitutes useful information in estimating potential employment and the employment gap in Canada. If the current level of the NAIRU is 8 per cent, say, the equilibrium employment-population ratio according to Table 3 is 59.9 per cent. This is just 0.4 of a percentage point higher than the actual mid-1998 employment-population ratio. Given our 24 million working-age population, the implied current employment gap would be only 96,000 jobs. If instead the current NAIRU is 6 per cent, the equilibrium employment-population ratio is 61.7 per cent. The resulting 2.2-percentage point differential with the actual mid-1998 employment-population ratio translates into a much larger employment gap of 528,000 jobs.

Finally, our statistical results are ill-suited for appraising future structural developments in labour force participation. There are three main reasons for this. First, our modelling of past non-policy structural change by equation 5 is based on mechanical annual time trends that are mute on the nature and relative contribution of underlying structural factors. Statistical tests reported above have confirmed the reliability of these trends for 1990-96, but projecting them blindly into the future would be foolhardy.

Second, assuming continuation of the past estimated trends reported in Table 1 leads to the absurd prediction that group-specific and aggregate labour force participation rates in Canada are going to fall at an accelerating rate in coming years. In particular, the participation rate of middle-aged women would be projected to drop from its current 77 per cent to the Stone Age level of 45 per cent before 2020. This does not make much sense.

Third, available data indicate that, contrary to the pattern of accelerating decline reflected in our estimated equations, all group-specific participation rates except for young men have stopped declining or begun to rise again since 1995. In this respect, the papers by Ip, King and Verdier (1999) and by Dugan and Robidoux (1999) in this symposium represent useful attempts to escape from pure projections of past trends. Their analysis based on systematic comparisons between Canada and the United States takes off where our econometric study ends.

Conclusion

The 2.7-point drop in its aggregate labour force participation rate that Canada suffered from 1989 to 1997 is both a puzzle and a cause for concern. The puzzle is that this occurrence was totally unexpected and unprecedented in the past half-century, and has remained mysterious to this day. The concern is that, if the drop in participation is not reversed or at least offset by some concomitant decline in the equilibrium unemployment rate, Canada will suffer from a permanent reduction of its national standard of living. There will be permanently less employment, production and real income per capita.

So, should the drop in Canadian labour force participation be a cause for concern? Yes and no. First, poor macroeconomic conditions in the 1990s have been unambiguously bad for participation and income, but this can be reversed if and when the Canadian economy recovers to full employment. Also worrisome is the high labour force dropout rate among the low-skilled, which can only be reduced by education, training or other long-term structural intervention. In addition, the welfare implications of other participation-reducing factors, such as minimum wage increases, restricted access to unemployment insurance benefits and more generous public pensions, are controversial. To decide whether those developments are good or bad requires a value judgment about the relative merits of their implications for efficiency and equity. To the extent that they result from free choice, the rising school enrolment rate and the stabilization of women's participation rate should be viewed as welfare-enhancing. In particular, it is worth pointing out that students who participate less in the labour force today are likely to participate more, and more productively, later.

We do not anticipate that aggregate labour force participation in Canada will return to its late-1980s level in the foreseeable future. Minimum wage, unemployment insurance, and social assistance policies seem to have reached a stable plateau, and any changes in pension policy will have effects only over the long run.

The current non-inflationary recovery from the slump of the last decade will lead to some increase in participation and employment. But even moderately optimistic estimates of the non-inflationary level of the unemployment rate (the

NAIRU), such as 6 per cent, would raise the aggregate participation rate only one third of the way back to where it was 10 years ago. If a non-inflationary decline of unemployment to 6 per cent indeed occurs, it would in fact offset much of the impact of the drop in participation on employment. We estimate that, within the current (1997) structural setting, a 6 per cent NAIRU would imply a 65.6 per cent aggregate participation rate, a 61.7 per cent employment-population ratio, and a remaining short-term employment gap of more than 500,000 jobs.

The type of statistical analysis in this paper is not well-suited for appraising future non-policy structural developments in labour force participation. Our study should be seen as complementary to other papers in this symposium that examine this problem with alternative methodologies.

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Notes

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1. We are aware that Statistics Canada's survey methodology, the composition of job offers, and the extent of labour market mismatches may have changed the meaning of the help-wanted index over the 28-year period studied. The statistical implication of this measurement error is that, in the regressions in first-differences reported below, the impact of the true change in job availability on the change in labour force participation will be somewhat underestimated. However, given the high simple correlation (90 per cent) between changes in the help-wanted index and changes in the employment-population ratio, we trust that the measurement signal-to-noise ratio is large and the estimation bias is therefore small.

2. One acceptable alternative is to rely on the change in the help-wanted index as instrument to exogenize the change in the employment-population ratio and use the projection to measure ΔJ and estimate equation 4 (in first-difference form). Introducing the change in real gross domestic product or the change in any measure of the output gap as second instrument adds no further explanatory power. We will employ this projection of the employment-population ratio in the simulations reported in the section looking to the future.

3. Wage relativities (e.g. women versus men, younger versus older, low-skilled versus high-skilled) also matter in principle for the labour force participation decision, but we do not have good macroeconomic measures of these variables. We assume that their influences are captured indirectly through other variables in our model equation 4.

4. In addition to the UI wage subsidy variable just defined, we experimented with two other UI program parameters: the national replacement rate (separately), and the ratio between maximum insurable and average earnings. Their estimated effects on participation were never statistically different from zero. We did not try to use the disqualification rate for quitters as a proxy for legal and administrative pressure on UI claimants, because this is a strongly endogenous variable for which there is currently no obvious instrument. We thank one of our referees for insisting that we test alternative UI variables.

5. This is shown by formal tests of unit roots based on the 'augmented Dickey-Fuller' statistical procedure.

6. In our particular application, where the regressors are the same in all equations, Zellner's method gives the same estimated coefficients as the ordinary-least-squares method, but smaller standard errors around these coefficients.

7. In view of the simultaneous feedback between labour force participation and wages, the equations including $\Delta\log(W/P)$ were estimated with the three-stage-least-squares method. Identifying instruments were past real wage growth, past CPI inflation, and a dummy variable meant to capture the productivity slowdown of the past 20 years. The joint hypothesis that the β_{i2} coefficients for $\Delta\log(W/P)$ estimated by this method are zero for all six groups cannot be rejected at standard levels of confidence. The probability value of the calculated chi-square statistic is 0.40.

8. We checked this in two ways. First, we added two variables to the list of regressors of the six equations: a dummy variable equal to 0 before 1990 and to 1 in 1990-96, and a new trend variable equal to 0 before 1990 and to 1, 2,...,7 in 1990-96. Each of the 12 estimated coefficients turns out to be statistically negligible. Second, we added a quadratic time trend term in each of the six difference equations, which amounts to introducing a cubic term in the level-form equations 4. Again, each of the six estimated coefficients turned out to be statistically negligible.

9. With given job availability and social policy, female participation rates stop rising when their associated $\Delta\log(L_i)$ declines to zero. This occurs in year $t_0 = 1969 - \beta_{i0}/\beta_{i6}$. For example, since Table 1 estimates for middle-aged women are $\beta_{i0} = 5.31$ percentage points and $\beta_{i6} = -0.190$ percentage point per year, their participation rate is estimated to become flat in year $t_0 = 1997$.

10. The contradiction with Dugan and Robidoux (1999), who find a much larger demographic composition ef-

fect, is only apparent. It occurs for two reasons. First, they use 1989 labour force weights instead of our average of 1989 and 1997 weights. Second, they break down the labour force into 16 demographic groups, whereas we use only six. As a result, part of their estimated demographic composition effect is captured by our estimated within-group trend effects.

11. Our simulations assume this SA benefit index remained unchanged in 1997 from 1996.

12. Beaudry and Lemieux (1999) present a finer analysis of the diffusion of labour force participation across successive cohorts of women in 1976-94.

13. In the absence of population shifts, equation 7 simply reads $\Delta L = \Sigma A_i \Delta L_i$, where the A_i's are the population shares. It follows that $\Delta\log(L) \approx \Delta L/L = \Sigma (A_i L_i/L)(\Delta L_i/L_i) \approx \Sigma F_i \Delta\log(L_i)$, since for each group i the expression $A_i L_i/L$ simplifies to its labour force share F_i. Hence the result $\beta_1 \approx \Sigma F_i \beta_{i1}$, that the cyclical elasticity of the aggregate participation rate (β_1) is approximately equal to the labour-force-weighted average of the cyclical elasticities of the group participation rates (β_{i1}).

14. With this replacement, we have estimated the six group equations by three-stage least squares to get around the simultaneous feedback between participation and employment. The identifying instruments used are last year's log change in E and the current growth rate in real GDP. This procedure gives statistical results almost identical to those reported in Table 1.

15. Achieving a 6 per cent unemployment rate in Canada may further require that the inflation rate be allowed to stabilize close to 3 per cent instead of around 1 per cent. Wage resistance and liquidity trap constraints on achieving the NAIRU have been emphasized recently by researchers such as Summers (1991), Tobin (1995), Akerlof, Dickens and Perry (1996), and Krugman (1998).

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Data Appendix

Labour force participation rates (L_i): Proportion of group population i (i = men 15 to 24, women 15 to 24, men 25 to 54, women 25 to 54, men 55 and over, women 55 and over) who have a job or who are actively looking for one. Data come from Statistics Canada's monthly Labour Force Survey and are extracted from the CANSIM database. For 1977-97, the log change in L_i is calculated from the most recent revised series for 1976-97; for 1969 -76, the previous series for 1968-76 is used.

Index of job availability (J): Ratio of help-wanted index to total population aged 15 and over. The help-wanted index comes from Statistics Canada's monthly measurement of newspaper job offers. We have put together a continuous series for 1968-97 by linking successive published revisions of the index. The population aged 15 and over comes from the monthly Labour Force Survey. All data are extracted from the CANSIM database.

Average real wage (W/P): Ratio of the average wage (W) to the consumer price index (P). For 1983-97, the average wage is the new series for the fixed-weight index of average hourly earnings of all employees; for 1968-82, it is the old series for the average weekly earnings of all industries. The two series overlap in the first three months of 1983 and are linked on that basis before the log change is calculated. Data come from Statistic Canada's monthly establishment Survey of Employment, Earnings & Hours. The consumer price index comes from Statistic Canada's monthly Consumer Price Survey. All series are extracted from the CANSIM database.

Average real minimum wage (W_m/P): Ratio of the average minimum wage (W_m) to the consumer price index (P). The average minimum wage is a fixed-weight average of individual provincial minimum wages, with weights equal to the 1991 provincial labour force shares. Data come from the Department of Human Resources Development.

Average UI wage subsidy (rD/M): Standardized product of the wage replacement rate (r) by the maximum duration of UI benefits available to the minimally qualified worker (D weeks), divided by the length of the minimum qualifying period (M weeks). The replacement rate is set at the statutory national value. In relevant years, it is a fixed-weight average of the general rate and the rate for claimants with family responsibilities. The replacement rate is also adjusted to reflect the non-taxability of UI benefits before the June 1971 UI Act. For the period before June 1971, the D/M ratio is a fixed-weight average of the nationally set regular and seasonal ratios. For the period after June 1971, M and D are national fixed-weight averages of standardized provincial values. These standardized values are those M and D would have taken under existing legislation with provincial unemployment rates fixed at their sample averages. For 1997, M is calculated under the assumption of a 30-hour standard work week. Data come from the various UI/EI Acts, and from Statistic Canada's quarterly reports on UI (cat. no. 71-001 and 73-001).

Average real SA benefit (B/P): Ratio of the average SA benefit (B) to the CPI (P). The average SA benefit is the ratio of the sum of all provincial expenditures on general social assistance to the corresponding total recipient population. For any given year, the latter is approximated by the average of current- and next-year March observations. Data come from the Department of Human Resources Development.

Time trend (T): Variable equal to 0 in 1969, 1 in 1970, 2 in 1971,..., 27 in 1996, 28 in 1997.