The Determinants of the Labour Force Participation Rate in Canada, 1968-97: Evidence from a Panel of Demographic Groups

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1. Introduction

Between 1990 and 1995, five successive years of decline in the Canadian participation rate pushed it down from 67.5% to 64.8%, a fraction which remained stable thereafter. Because this sudden and significant decline in the trend participation rate has no precedent, it has become, with the sustained high level of unemployment, the major fact of Canada's labour market in the 90s. Since the sudden occurrence of the decline and its persistence call for diverging explanations, a single cause is not likely. While abrupt movements naturally leads to suspect cyclical factors, the persistence generally points toward structural explanations. Given that the 90s have witnessed the most prolonged cyclical slump in Canada since the 30s, there is no ambiguity as to whether or not the usual pro-cyclical reaction of the participation rate may have played a role in the decline. Yet, the relative contribution of cyclical and structural factors are essential to address correctly the issue. If it were mainly a cyclical decline, then the participation rate could return close to its previous high of 67.5%, which implies that it remains presently a significant output gap. At the opposite, potential output would be close to actual output if structural factors explain much of the decline.

This paper disentangles the change in the participation rate according to three types of factors, that is, cyclical, trend, and social programs. Moreover, since the repartition between these factors varies with age and sex, we analyze the movements of the labour force along three age groups (15-24, 25-54 and 55 and over) for both sexes. This exercice is an update and a complement of Fortin and Fortin (1997) in which we estimated the contribution of these same factors for a panel of five regions and six demographic groups. Despite that we then found significant statistical differences in the regional behaviour of females,[1] we will not focus on these regional discrepancies in this study to concentrate our attention to the more important differences observed amongst demographic groups.

[1] Our last year study reveals also that young males behaved slightly differently in the Atlantic region than in the rest of the country.

Some of the demographic groups we havedefined have received a particular attention recently. Grignon and Archambault (1998) analyze how structural and cyclical factors explain the decline in the youth participation rate with a particular emphasis in distinguishing between students and non-students. Lemieux and Beaudry (1998) analyze more specifically the role of cohort in explaining the changing trend of females. Our contribution in this literature in twofold. First, by applying the same model to all demographic groups, we analyze in a consistent way the participation rate. Second, we put a particular emphasis in the measure of the impact of legislative changes in the unemployment insurance (UI) program. This emphasis on UI leads us to make three fondamental decisions regarding the appropriate treatment of this variable.

First, we think it is very important to use a sample that begins in the 60s. As many before have pointed out, the 1971 reform of UI is the main candidate to explain why the Canadian unemployment rate has not been declining with respect to the U. S. rate in the 70s despite the fact that the employment ratio was growing much faster in Canada (See, for example, Card and Riddell, 1993). Indeed, one of the main prediction of the classical labor supply is that when UI heavily subsidizes sporadic employment, this will makes some marginally attached workers to move in the labour force (Fortin, 1984). Following the 1971 reform, a subsidy rate as high as 367% was possible in some regions of Canada. Subsequent changes somewhat reduced this maximum subsidy rate but it was mainly in the 90s, with a succession of legislative changes in 1990, 1994 and 1996, that the generosity of the program was significantly reduced to return closer to the level of generosity seen before the 1971 reform of US.[23] Given that the participation rate reaction of the 90s to UI is just the reverse reaction of the 70s, it becomes important that the empirical model is able to explain as well the rise of the 70s than the fall of the 90s.

[2] The concept of generosity we use if the maximum subsidy rate of employment income by unemployment insurance. Our calculation show that the Canadian average of this subsidy rate was approximately 31.5% in the 60s, reached approximately 2..5 between 1972 and 1975, and was still 1.8 in 1989. Since the last three legislative changes of the 90s, the subsidy rate has fallon to approximately 0.62.

Second, we do not directly use the generosity of the program but rather an instrument for this variable. The reason behind this decision lies in the retroaction between the participation rate and the generosity of the program. One of the characteristics of the UI program is that the maximum duration of benefits and the qualification condition are dependent on the regional unemployment rate. If for some reasons the participation rate changes, this implies a response of the unemployment rate which, in turns, influences the generosity of the program and the explanatory variable. This violates the assumption of independence between the error terms and the regressors. To avoid a biased measure of the impact of UI, we calculated an index of the generosity of the program that responds solely to legislative changes. Finally, our third contribution lies on the explicit recognition that the response of the participation rate to UI changes may not be immediate. There are empirical elements, already pointed out namely by Fortin (1994) and by Archambault and Fortin (1997), which strongly suggests that this could be the case.[4] Thus, we estimate a lag structure for the UI variable in the participation rate equation.

The paper is organized as follows. The next section presents a theoretical model of labour force participation. Data and estimation of a system of equations for six demographic groups are presented in section 3. The fourth section indicates the implications of the results for a global participation rate equation and compares with an estimation of a single participation rate equation. Concluding remarks close the paper in the last section.

2. The theory

We consider an individual who must choose between entering into the labour force, which may lead to employment or unemployment, or remaining out of it and be inoccupied. The monetary value of the reservation value of time of the individual *i* at period *t* is given by $\text{"}_{it}(\mathrm{J}_t)$, where J_t is an

[4] The bottom ligne of these elements lies on the simple observation that between 1971 and 1978, the Canadian Beveridge curve drifts on the right at a decreasing speed.. Although by no means a proof that the 1971 reform is solely responsible for this movement, is is certainly consistent with a gradual response of the participation rate.

index of its non-labour income. When deciding to be active, it is uncertain if it will occupy a job or remain unemployed. The expected payoff from participating is then an average of the income received while unemployed or employed. In the former case, he must be searching for a job and pay a cost *c*, *w*hich includes both direct costs and indirect inconvenients related with searching, so that its satisfaction level is $"_{it}(\mathrm{J}_t)-c$. As to employment, we assume that at period *t*, all jobs pay an identical wage w_t. However, because of UI, wage is not the only component of remuneration. Given that an employee works at least m_t periods, it is entitled to receive d_t periods of UI benefits set at a fraction R_t of the wage. Following Fortin (1984), we define the maximum subsidy rate D_t as the product of the benefit rate R by the ratio d_t/m_t, that is, $\mathrm{D}_t = \mathrm{R}_t \times d_t/m_t$. In addition to the labour income w_t, each period of work thus entitles the worker to UI benefits of a value $\mathrm{D}_t w_t$. By assuming that benefits rights are valued as much as labour income, the total income received for each period of work is then $w_t(1+\mathrm{D}_t)$.[5]

We complete the description of the expected payoff by a description of the probability $p(\mathrm{N}_t)$ of beeing employed if someone participates to the labour market. We suppose that N_t is an index of business cycle conditions such that $dp(\mathrm{N}_t)/d\mathrm{N}_t > 0$. The expected payoff of participating to the labour market is then $p(\mathrm{N}_t)w_t(1+\mathrm{D}_t) + [1-p(\mathrm{N}_t)]["_{it}(\mathrm{J}_t)-c]$. Individual *i* participates to the labour market if the following inequality is verified :

$$p(\mathrm{N}_t)w_t(1+\mathrm{D}_t) + [1-p(\mathrm{N}_t)]["_{it}(\mathrm{J}_t)-c] \ \$ \ "_{it}(\mathrm{J}_t) \qquad (1)$$

The marginal active person on the labour market satisfies equation (1) with equality. If $"_{it}^{*}$ is the reservation value of time of this marginal person, we then have the equality $p(\mathrm{N}_t)w_t(1+\mathrm{D}_t) + [1-p(\mathrm{N}_t)]["_{it}^{*}(\mathrm{J}_t)-c] = "_{it}^{*}(\mathrm{J}_t)$. With a simple manipulation, one can write :

[5] We must emphasize that *d* is not the maximum duration of benefits but the maximum number of benefits weeks for a minimally qualified person. As in Fortin (1984), the main determinant of the change in the participation rate is the maximum subsidy of employment rate.

$$"_{it}^{*}\,(\mathrm{J}_t) = w_t(1+\mathrm{D}_t) + c[1\text{-}p(\mathrm{N}_t)]/p(\mathrm{N}_t) \tag{2}$$

As a last step, we analyze the aggregate consequences of these individual decisions. Let us assume that $"_{it}$ follows the marginal continuous distribution f($"_{it}$). The cumulative distribution is then F($"_{it}$). Since the labour force A_t is equal to the number of persons that satisfies equation (1), is is thus equal to F($"_{it}^{*}$), that is, the cumulative distribution evaluated to the marginally attached person. The labour force equation is then given by :

$$A_t = \mathrm{F}("_{it}^{*}(\mathrm{J}_t)) = \mathrm{G}(w_t(1+\mathrm{D}_t) + c[1\text{-}p(\mathrm{N}_t)]/p(\mathrm{N}_t),\mathrm{J}_t) \tag{3}$$

Dividing by the population of 15 and over (P_t) gives the participation rate $a_t = A_t\,/P_t$. Let g($\cdot$) = G($\cdot$)/P_t. We can then write :

$$a_t = \mathrm{F}("_{it}^{*}(\mathrm{J}_t)) = \mathrm{g}(w_t(1+\mathrm{D}_t) + c[1\text{-}p(\mathrm{N}_t)]/p(\mathrm{N}_t),\mathrm{J}_t) \tag{4}$$

It is easy to show that a_t increases with $w_t(1+\mathrm{D}_t)$ and with N_t but decreases with J_t. This implies that the participation rate is pro-cyclical and an increasing function of the maximum subsidy rate of employment income by unemployment insurance. However, the activity rate is negatively affected by the non-labour income J_t. These last two predictions are consistent with the model presented in Arnau, Crémieux and Fortin (1998). Finally, the model predicts that the wage unambiguously raises the participation rate.

Before turning to estimation, some remarks help in translating the theoretical predictions into real world expected impacts. First, the predicted positive impact of the real wage is not consistent with respect to the usual ambiguous impact of wages on labour supply. The reason for this difference lies in the implicit assumption that individuals can work either full time or not at all. Because of these limitations, the budget constraint is composed of only two points in the expected income/time space.

A rise in wage improves the expected payoff from participating while leaving unchanged the satisfaction level of non-participation. In a less rigid environment in which individuals can choose the fraction of working time, the positive income effects of a higher wage allows individuals to work a smaller fraction of their total time. This could be achieved either by working for a shorter time each period, or by reducing the number of periods working full time. In the second case, the aggregate consequence would be that the participation rate could fall when the wage rate rises. Since we believe that this reaction is possible, we make no a priori assumption regarding the impact of wage.

A second remark is in order concerning the minimum wage. It has long been recognized that the minimum wage has two opposite consequences on labour market participation. Because better paid jobs raise the expected payoff, more minimally qualified workers may be drawn into the labour force than would be the case in the absence of a minimum wage. However, these higher wages may reduce the number of jobs available. Therefore, as the model shows, a lower probability of finding a job depresses the participation rate. Thus, and despite the fact that the model explicitly takes into account the wage impact and the employment probability, it is possible that the empirical variables do not capture the specific impact of the minimum wage on participation. Since this impact depends of the effect of the minimum wage at a given distribution of wages, we add as an additional variable the ratio of the minimum wage to average wage. The theoretical participation rate equation is then of the form :

$$a_{it} = f(N_t, w_t, wmin_t/w_t, D_t, J_t) \qquad (5)$$

with an expected positive impact of N_t and D_t, and a negative influence expected for the non-labour income J_t. As to w_t and $wmin_t/w_t$, they can be of either sign.

3. The empirical model and the data set

The participation rate of the entire working age population is an average of the participation rate of many demographic groups which sometimes have very different behaviour. This is easily illustrated by the fact that the trend in the participation rates of males and females have been of opposite signs over the last thirty years. This has for consequence that when group shares are changing, the aggregate pamameters becomes unstable and, because of this, may lead ot unconsistent estimates. To circumvent the problem, we estimate the parameters for six demographic group, that is, male and female for the age groups 15-24, 25-54 and 55 and over. From these estimates, we recover the global equation by simply calculating the group average for each coefficient. Since the participation rate in level is non-stationnary, the following system of equations was estimated in first difference :

$$\Delta \log(a_{it}) = \beta_{i0} + \beta_{i1} \Delta \log(N_t) + \beta_{i2} \Delta \log(w_t) + \beta_{i3} \Delta \log(wmin_t/w_t) + \beta_{i4} \Delta \log(SAB_t/wmin_t)$$
$$+ \beta_{i5} \Delta \log(UIG_t) + \beta_6 \Delta T_t + u_{it,} , \qquad i = 1, \ldots, 6 \tag{5}$$

The left hand variable is the logarithmic change in the participation rate of group *i*. All right hand variables, with the sole exception of the time trend, are also expressed in logarithmic differences. The first variable is a cyclical indicator of the probability of finding a job. A natural choice for this cyclical indicator is the employment/population ratio(e_{it}). However, because some of the movements in e_{it} are the results of changes in a_{it}, using directly e_{it} as a regressor in the equation would create a simultaneity bias. More precisely, the coefficient of e_{it} would over-estimate the true cyclical sensitivity of the participation rate. In order to avoid this problem, we could as in Fortin and Fortin (1997), use a two stage least square to estimate an instrument for $e_{it.}$. We rather preferred the simpler method of instrumenting the employment/population ratio with the ratio of the help-wanted index (HWI) to the population of 15 and over. Archambault and Fortin (1997) showed that an index of the vacancy rate based on the HWI reacts strongly and fastly to cyclical shocks but is unsensitive to shocks to the particpation rate,[6] which makes it a good instrument for capturing the cyclical movements in the probability of finding a job. We divided the HWI by the population to correct for the growing size of the labour force.

The second variable is the logarithmic change in the real consumption wage, that is, the nominal wage divided by the consumer price index. The third variable is the logarithmic difference in the relative minimum wage, which is defined as the ratio of the minimum real wage to the average real wage. The fourth variable is the logarithmic change in the relative social assistance benefit (SAB), which is defined as the ratio between, on one hand, total benefits in real value divided by the number of beneficiaries and, on the other hand, the real minimum wage. As Arnaud, Crémieux and Fortin (1998) show, the real social assistance benefits, which is the minimum out of work income, influences working incentives because it reduces the marginal net income from working. Since social assistance benefits are mainly used by low skilled workers whose main labour market alternative is non-specialized work paid a the minimum wage, this ratio crude measure of the potential impact SAB can have on labour market participation is the logarithmic difference between real social assistance and the minimum wage.

The fifth variable captures the impact of UI benefits on the participation rate. We use $\log(1+D_t)$ because in the theoretical model, employment income is defined as $w_t \times (1+D_t)$ so that this product is additive in the logarithm. The UI subsidy rate is defined as the average of the subsidy rate in all regions of Canada. Moreover, to eliminate the retroaction between the unemployment rate and the index of the generosity of UI, we use a standardized index calculated not on the actual unemployment rate but under the hypothesis that the unemployment rate has remained constant to

[6] The conclusion was based on the fact that in a VAR, the variance decomposition of the HWI into cyclical, structural and participation rate finds a contribution of cyclical shocks close to 100% at any time horizon. We further checked the appropriateness of this instrument by regressing the logarithmic change in the employment population ratio on the logarithmic change of the instrument and on the logarithmic change of the real per capita GDP over the period 1967-1997 and only the first variable stands out as a significant regressor. Morevoer, the adjusted R^2 of this regression (0.735) is lower than when the employment population ratio is explained solely by the instrument (0.742).

its sample mean. Thias implies that the variable is an instrument that reflects solely changes in legislation. Finally, because the labour supply reaction to UI changes is likely to be progressive, we used a moving average of current and lagged values of $(1+D_t)$. The length of this MA has been chosen to maximize the adjusted R^2 and to minimize the information criteria. Finally, the last variable is a simple time trend. Since the model is estimated in first difference, the trend growth rate of the participation rate is captured by the constant. The time trend allows this growth rate to change progressively. Because the variables are in log rate of change, each coefficient has the dimension of an elasticity.

The data set is composed of annual observations covering the period from 1966 to 1997 for all the variables with two exceptions. The data for the social assistance benefits were available only from 1968 to 1996. As to the index of UI generosity, it was easy to calculate its value back to the beginning of the sixties (See appendix). Because the series are in first difference, the estimation is made on 28 annual data from 1969 to 1996 for 6 groups, that is, a total of 168 observations. We estimated the system of equations by the method of Zellner's Seemingly Unrelated Regression (SUR). Since we use the same set of regressors in each equation, SUR estimates are identical to OLS estimates. Yet, SUR estimates are nevertheless more efficient thant OLS. By taking account of the correlation of the errors between equations, SUR provides smaller estimated variances and more powerful tests. We used the fact that the estimates were identical to OLS to explore the impact of various specifications of the lag length of $(1+D_t)$ in each equation. It became rapidly obvious, as we will show in a short time, that UI has a significant impact particularly for young people. The information criteria from the OLS equation for young males and females suggest to keep three lags of $(1+D)$ in addition to the current value. Thus we retained the OLS estimates of the three lags of $(1+D)$ to calculate the MA.[7] Appendix 1 gives the source of all data.

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[7] The UI variable used in the model is) UIG = -0.09057×) $(1+D_t)$ + 1.14718×) $(1+D_{t-1})$ + 0.18548×) $(1+D_{t-2})$ - 0.24021×) $(1+D_{t-3})$. We constraint the sum of the weights to be one to preserve the interpretation that the coefficient is an elasticity. In subsequent steps of this study, we will explore more thorouhgly the possibility of using different lag structures in each equation.

3. The estimation results

The first set of results are those of the unconstrained SUR, reported in table 1. These equations have been submitted to many tests to detect any symptom of misspecification. The LM test with two lags do not detect any significant autocorrelations of the residuals. The RESET test do not allow to reject the null hypothesis that one or two fitted terms have no impact on the dependent variable. The stability of the parameters was checked by the means of Chow tests with a breakpoint in 1981 and by looking at recursive estimates. Once again, ne problems were detected.

Table 1
Unconstrained SUR estimates

Variables	15-24 (M)	15-24 (F)	25-54 (M)	25-54 (F)	55+ (M)	55+(F)
Constant	0.0092* (0.0054)	0.0311^^ (0.0069)	0.0018 (0.0015)	0.0593^^ (0.0044)	-0.0034 (0.0096)	0.0322^ (0.0145)
N_t	0.0654^^ (0.0074)	0.0423^^ (0.0093)	0.0130^^ (0.0021)	0.0099* (0.0059)	-0.0021 (0.0130)	0.0034 (0.0196)
w_t	-0.1042 (0.0911)	-0.2562^ (0.1143)	-0.0603^ (0.0253)	-0.2053^^ (0.0730)	-0.1355 (0.1599)	-0.3288 (0.2408)
$wmin_t/w_t$	-0.1618^^ (0.0553)	-0.1587^ (0.0693)	-0.0058 (0.0154)	-0.1556^^ (0.0443)	-0.1241 (-0.1197)	0.1508 (0.1460)
SAB_t	0.0146 (0.0371)	-0.0342 (0.0465)	-0.0112 (0.0103)	-0.1041^^ (0.0297)	-0.1197* (0.0650)	-0.1709* (0.0979)
UIG_t	0.0703^^ (0.0167)	0.0571^^ (0.0209	-0.0017 (0.0046)	0.0148 (0.0133)	0.0126 (0.0292)	-0.0964^ (0.0440)
T_t	-0.0005* (0.0003)	-0.0013^^ (0.0003)	-0.00015^ (0.00007)	-0.0019^^ (0.0002)	-0.0007 (0.0005)	-0.0017^ (0.0007)
R^2	0.8632	0.7892	0.6866	0.8534	0.1680	0.2726
Adj. R^2	0.8241	0.7289	0.5970	0.8116	-0.0698	0.0648
S. E.	0.0082	0.0103	0.0023	0.0066	0.0144	0.0217
D. W.	1.8023	2.2407	2.5210	2.0759	2.6336	3.0806

The symbols *, ^ and ^^ indicate that the coefficient is significant at the 10%, 5% and 1% level respectively. Standard deviations are between parenthesis.

The R^2 of the model varies considerably in the different equations. The model performs very poorly in explaining the time series behaviour of the participation rate of both males and females of 55 and over. This indicates that the changes in the participation rate for older people is not sensitive to the business cycle and there is no influence of social programs at the usual statistical level. However, the equations explain more than 80% of the variations in the participation rate of males 15-24 and of females 25-54 and close to 80% for females 15-24. This fraction is a bit smaller for male 25-54 but reaches nevertheless 68%. Given that the estimation is in first difference and that the lagged value of the dependent variable do not appear in the specification, this denotes a high explanatory power of the model. We now turn our attention to the impact of each variable.

i. Cyclical variable

The model shows large differences in the cyclical sensitivity of the participation rate in the various demographic groups. The greatest sensitivity is observed for young males, with an elasticity of the participation rate to N_t of 6,5% while it is only 1,3%, that is, five times smaller in the age group 25-54. As to women, the elasticity is about 2/3 the elasticity observed in males for the same age group, and the cyclical variable is significant only at the 10% level in the equation for women 25-54. As for older people, the model do not detect any statistically significant response of the participation rate to cyclical changes.

ii. Real wage

The coefficient of the real wage is negative in all six equations and is significant at the 1% level for females 25-54 while it is significant at the 5% level in the equation for males 25-54 and for females 15-24. However, since the minimum wage is expressed in log differences with respect to the real wage, the correct impact of the real wage is measured by the difference between the coeffcient of these two variables.[8] The null hypothesis that the real wage has no impact in all equations has a P^2 of 6.304990 and a marginal probability of 0.504623. Thus, one cannot reject the null hypothesis that the real wage plays no role in explaining the participation rate for all demographic groups.

iii. Relative minimum wage

Since the minimum wage divides the social assistance benefits, its impact is measured as the difference between the third and the fourth coefficient in each equation. This variable has a negative impact on the participation rate for the youth, for femal 25-54 and for males 55+. Its impact of the participation rate is significantly lower than zero at the 1% level for young males and at the 5% level for young females. It is significantly positive at the marginal level of 1.3% for female workers 55+.

iv. Relative social assistance benefits

The relative social assistance benefits has a negative impact for five groups. However, the usual 5% significance level is attained only for women 25-54. For both men and women over 55, the impact is significant at the 10% level and the largest coefficient (-0.17) is found for women over 55.

v. Unemployment insurance generosity

For both males and females 15-24, this variable has a similar positive impact (0.07 for males and 0.057 for females) that is significant at the 1% level. The model also detects a significant impact at the 5% level for females over 55 but the sign is negative. As to males and females 25-54, the model does not detect a significant reaction to the UI generosity.

vi. Time trend

The trend growth rate of a_t is captured by the constant while the time trend indicates if the trend growth rate is declining of increasing. For females of all age, the growth rate was initially

[8] Since the estimated equation is involves the expression $\$i_2$) log(w_t) + $\$i_3$) log($wmin_t/w_t$), this can be written as $\$i_2$) log(w_t) + $\$i_3$) log($wmin_t$) - $\$i_3$) log(w_t) = ($\$i_2$-$\$i_3$) log(w_t) + $\$i_3$) log($wmin_t$).

positive. However it was also declining so that at the end of the 80s, the trend became negative for 15-24 and 55+ age groups. For women 25-54, the trend becomes negative only in 1997. That means that during much of the 90s, the trend participation rate for middle aged women continues to increase to reach a maximum of almost 79%, three percentage points over the actual participation rate. As to males, the constant is never significantly different from zero although there is a significant decline in the growth rate for 15-24 and 25-54 age groups.

Table 2
OLS estimates for the global participation rate

Variable	Coefficient	Std. Error	t-Statistic	Prob
C	0.017191	0.002573	6.681813	0.0000
DLHWI	0.020115	0.003482	5.776331	0.0000
DWR	-0.155252	0.042740	-3.632503	0.0016
DWMR	-0.078762	0.025919	-3.038755	0.0062
DLPASR	-0.042268	0.017384	-2.431424	0.0241
XJH	0.019047	0.007811	2.438391	0.0237
T	-0.000652	0.000124	-5.252149	0.0000

Global statistics of the regression

R-squared	0.870099	Mean dependent var	0.004008
Adjusted R-squared	0.832984	S.D. dependent var	0.008150
S.E. of regression	0.003331	Durbin-Watson stat	1.491161
F-statistic	23.44354	Prob(F-statistic)	0.000000

To provide a basis for comparison, we present the aggregate model of the participation rate with the same specification. The dependent variable is the growth rate of the participation rate for 15 and over for both sexes. The R^2 of 0.87 is higher than for any equation estimated for the groups. Once again, this equation was analyzed for possible misspecification but the investigation did not detect autocorrelated residuals or unstable coefficients. We must indicate here that despite the high t-statistics for all variables, the impact of the real wage is given by the difference between the third and the fourth coefficent (-.0765), which difference has a marginal value of only 0.182 and is therefore not statistically significant. As to the minimum wage, its impact is given by the difference between the fourth and the fifth coefficient (-0.0365) and is significant only at the 0.11 marginal level

4. Decomposing the participation rate from 1989 to 1997

We can now calculate the contribution of cyclical and structural factors in the falling participation rate of the 90s. The empirical model explains the changes in the participation rate. We used the estimated coefficients to calculate the impact of each explanatory variables in the changes in the participation. Then, we computed recursively how this has affected the level of the participation rate. The level of all series has been set to be identical to the actual series for the year 1989. Thus, all subsequent results can be interpreted as the contribution of each variable to the evolution of the participation rate, if all other variables had kept their 1989 level.

Our first interest is in explaining the global participation rate. There are two ways to calculate the empirical impact of the variables. The first possibility is to use the estimated impact for each group to calculate an average for the whole population by using each group's population share. The second way is simply to use the global equation. As we discussed previously, the first method is preferable and was chosen for our first analysis of the global behaviour. Since the equations were estimated up to 1996, because we did not have tha 1997 data for social assistance benefits, the response for 1997 in an out of sample forecast

The figures 1 and 2 show the estimated response of the global participation rate to the explanatory variables, calculated from the results for the 6 groups. The fall of the participation rate between 1990 and 1997 was 2,7 percentage points. Five out of six factors have contributed to the fall, the only exception beeing a small positive impact (+0.3) of social assistance benefit. However, two variables plays a more important role in the fall of the 90s, that is, the trend and the cycle with a respective contribution of -1.24 and -1.03 respectively. The impact of UI is estimated to only -0.42

and is similar to that of the minimum wage (-0.42) and of the real wage (-0.50). These figures also show that between 1989 and 1991, almost all the fall in the participation rate was the result of cyclical factors. The business cycle continued to be the most important single cause of the decline in 1992 but since 1993 however, it is mainly structural factors that push down the participation rate with the trend as the dominant factor for 1996 and 1997.

The group analysis seems to be of importance to disentangle the sources of the change in the participation rate. We compare in figure 3 the responses to cyclical, trend and UI variables calculated from the global equation with those based on groups estimations. Our preferred method gives an impact of the trend significantly smaller than if we were using the global estimation. From 1989 to 1997, the estimates on the whole population imply that the trend would have reduced the participation rate by only 0.4 percentage points rather than 1.24 as we estimate from the groups estimates. As to UI, global results find a larger impact than when we calculate it from the groups, that is, -0.64 rather than -0.42. This is so because the youth, which are those with the greatest reaction to UI changes, represent a smaller share of the population, less than 17% in 1997. It is particularly interesting to note that the reverse was true during the 70s while the youth represented a share of more thant 26% of the population. Thus, the rise in the participation rate following the 1971 reform was magnified by the fact the more responsive groups were more important at this time. On the contrary, the response of the fall of the global participation rate to the restrictions to UI during the 90s is weaken by the low fraction of highly sensitive groups. We observe however that the cyclical fall in the participation rate remains very similar with both methods.

We now carry out a similar exercice for the groups. Because the estimated impact of the real wage, the minimum wage and of social assistance benefits is less important, we present in figure 4 only the trend, the cycle and the UI impact on the participation rate of males 15-24. The single most important source of the 10 percentage points drop observed between 1989 to 1997 is the business cycle, which caused a 3.5 percentage points decline. The second most important cause of the fall is the trend (-2.8) followed by UI, which caused a reduction of 2.5 percentage points. As in the case of the global participation rate, almost all the reduction in *a* that happened for young males between 1989 and 1991 is cyclical. Since 1992, cyclical factors have been replaced by the trend and UI as the main source of decline. As to females 15-24, the most important source of the 8.8 percentage points decline is the trend (-3.0) followed by the business cycle(-2.2) and UI (-1.9). Thus, in both young men and women, cyclical factors have been the source of almost all the decline in the beginning of the 90s. Since 1992 however, the downward movement observer to the young is mainly the consequence of reductions in UI and to the continuation of a long trend. The differences in the responsibility of the decline between males and females are then minor.[9]

For middle aged men and women, the decomposition is much different. For males, only the cycle and the trend had a significant impact on the participation. These two factors have been responsible for, respectively, 1 and 1.8 percentage point to the total decline of 2.7 percentage points. The contribution of all other factors is negligible. It is clearly the trend which is responsible for the decline since 1992.

The last figure shows the contribution of these same variables to the changes in the participation rate of women 25-54. Because of the trend, the participation rate continued to rise until 1991 despite the fact that the beginning of the recession started to push down the participation rate. Between 1989 and 1997, the participation rate increase by 0.9 percentage point, thanks to the trend which made for a 4 percentage point rise. The impact of cyclical factors were more limited on this group since the cyclical reduction was only 0.8 percentage point. In this group, both the real wage (-0.7) and the miminum wage (-1.0) reduced significantly the participation rate, by as much as cyclical factors.

[9] Because we concentrate our discussion towards explaining the 90s, we do not show the impact of UI following the 1971 reform. The model finds that at the aggregate level, the participation rate increased by approximately 0.9 percentage points between 1971 and 1975. For young males, the impact was almost five times larger, with a rise of 4.6 percerntage points during the same period while the estimated impact was 2.9 percentage points for young women.

We do not present any decomposition for older people because of the poor performance of the model for this age group. The F(6, 29) of the regression for men 55+ is only 0.70 with a marginal significance level of 0.65. As to women 55+, the F(6,29) = 1.31 with a marginal significance of 0.30. Thus, in both cases, one cannot reject the null hypothesis that the participation rate is driven solely by a constant time trend. Thus, even if individually the model finds a significant negative impact of social assistance and UI benefits on the participation rate of women 55+, the same model cannot help to explain the deviations of the growth rate of the participation rate from its long term mean.

5. Conclusion

It is crucial to understand if the declining participation rate was a response to demographic causes, the results of the macroeconomic disaster of the 90s or the consequence of changes in social programs. Our main objective was then to determine how the fall in the participation rate since the beginning of the 90s can be related to these causes. We have found that until 1991, the deteriorating macroeconomic conditions has been the main explanations to the decline. Since 1992 however, the falling participation rate is mainly due to a change in demographic. The changes in UI that have been introduced since 1990 have also contributed to the decline, but the impact is estimated to be less than half a percentage point, that is, a much less important source of fluctuations than the cycle or the trend. One of the explanation for this relatively small impact is the fact that the the youth are not numerous enough to weigth much on the global behaviour of the participation rate.

The repartition of cyclical and structural factors is very different among demographic groups. For people of 55 and more, the participation rate is driven only by a linear trend.Young workers react much more to cyclical fluctuations and toUI than middle aged workers. We found that for young males, macroeconomic factors have been the dominant cause of the declining participation rate, followed by UI and by the trend. As to middle age men, the only source of the decline since 1992 is the trend which implies a fall of approximately half a percentage point in the participation rate each year. We also found that the trend for women 25-54 is now zero.

We started this paper by asking what was the cyclical part of the declining participation rate? Our results show that if aggregate demand was returning to the same cyclical high as 1989, the participation rate would regain no more thant 1 percentage point, that is, about one third of the fall. The time when the trend of the global participation rate was beeing pulled up by the increased participation rate of middle aged women is now behind us. Since the participation rate of this group is no longer growing, the global participation rate is now driven by the declining participation rate of middle aged men and of older workers. This conclusion is shared by other researchers.

A natural complement to this work is to ask how our results impact on the potential output. To answer this question needs to complement our explanation of the participation rate by an analysis of the unemployment rate, or equivalently the employment ratio. With information regarding the productivity of the factors, we could calculate the size of potential output during the 90s. This is on our research agenda.

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Appendix : Source of the data

Participation rate and population : Statistic Canada. The data are based on two different estimates. For the period 1976-97, we used the participation rate adjusted to Census Data. The changes in the participation rate for the period 1966-76 were calculated from the unadjusted series.

Price : Consumer Price Index, Statistics Canada.

Wage : Statistic Canada. Since 1983, wage is the average weekly wage for all industries. Before 1983, we used the average weekly wage in manufacturing. These data are published in various issues of publications 72-002 and 72-202. Data prior 1983 have been adjusted to be consistent the new series on the first three months of 1983. The minimum wage is a national average of the provincial minimum wages. The nomimal average wage and the minimum wage have been divided by the CPI to obtain series in real terms.

Unemployment Insurance : Authors' calculation. Before the 1971 reform, there was regular ans seasonal benefits. To be entitled to regular benefits, it was necessary to work two weeks to be entitled to one week of benefits at a rate which varied according to the personal situation and the wage. We assume a gross replacement rate of 0.5, close to the fraction of 0.48 calculated by Sargent (1996).. As to seasonal benefits, which were paid at a rate identical to that of regular benefits, they represented approximately 35% of all benefits. One was entitled to 13 weeks of benefits for 15 weeks of work, or to a ratio of benefits weeks to working weeks of 5/6. For reasons that are exposed in Fortin and Fortin (1997), we do not agree with Sargent (1996) as to the possibility that in the sixties, it was possible to receive two weeks of benefits for each working week. Finally, benefits were not taxable so that an adjustment had to be made for the tax advantage. Our final estimate of the maximum subsidy rate is 0.315, which remained constant throughout the 60s.

Social assistance : Administrative data from Human Resources Development Canada. We defined the social assistance benefits as the ratio between total benefits paid and the average number of benefits at the beginning and at the end of the year. This variable was then divided by the CPI to obtain a real variable, and the result once again divided by the minimum real wage.

Appendix 2

Figure 1
Decomposition of A/P
based on groups

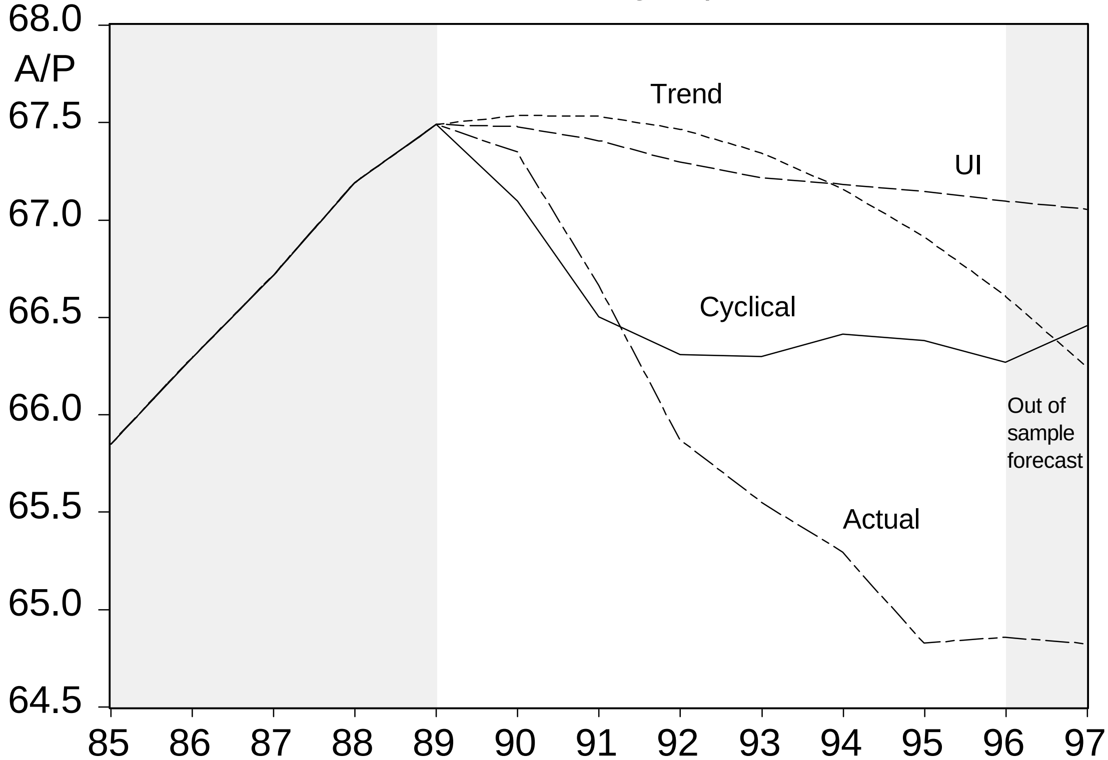

Figure 2
Decomposition of A/P
based on groups

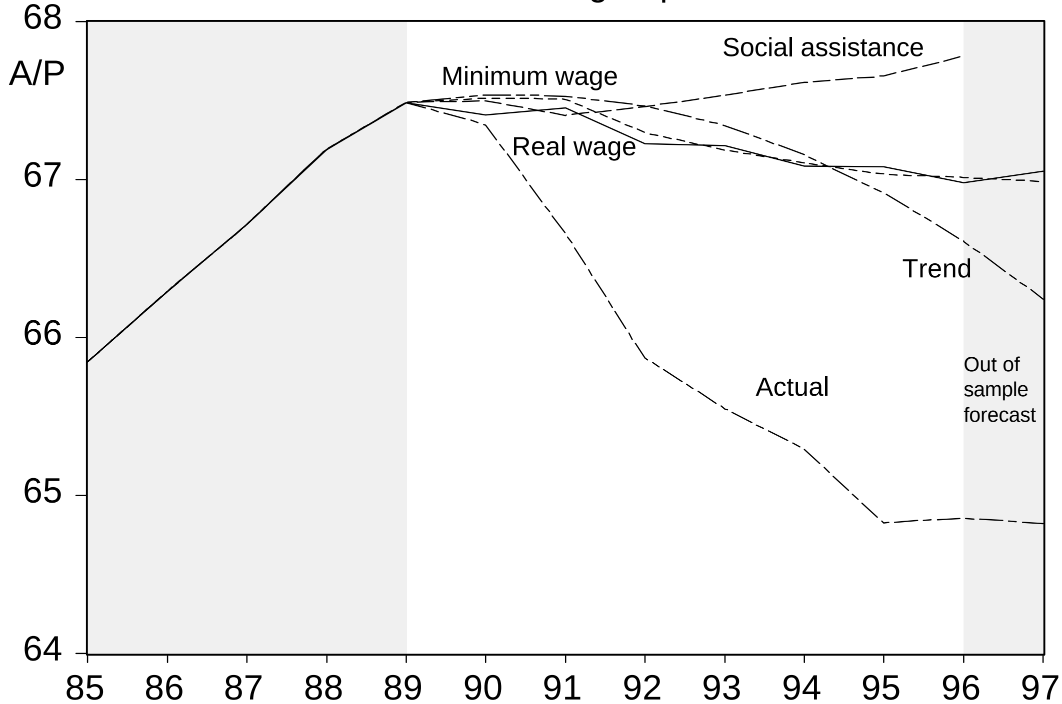

Figure 3
Comparisons of
global and groups responses

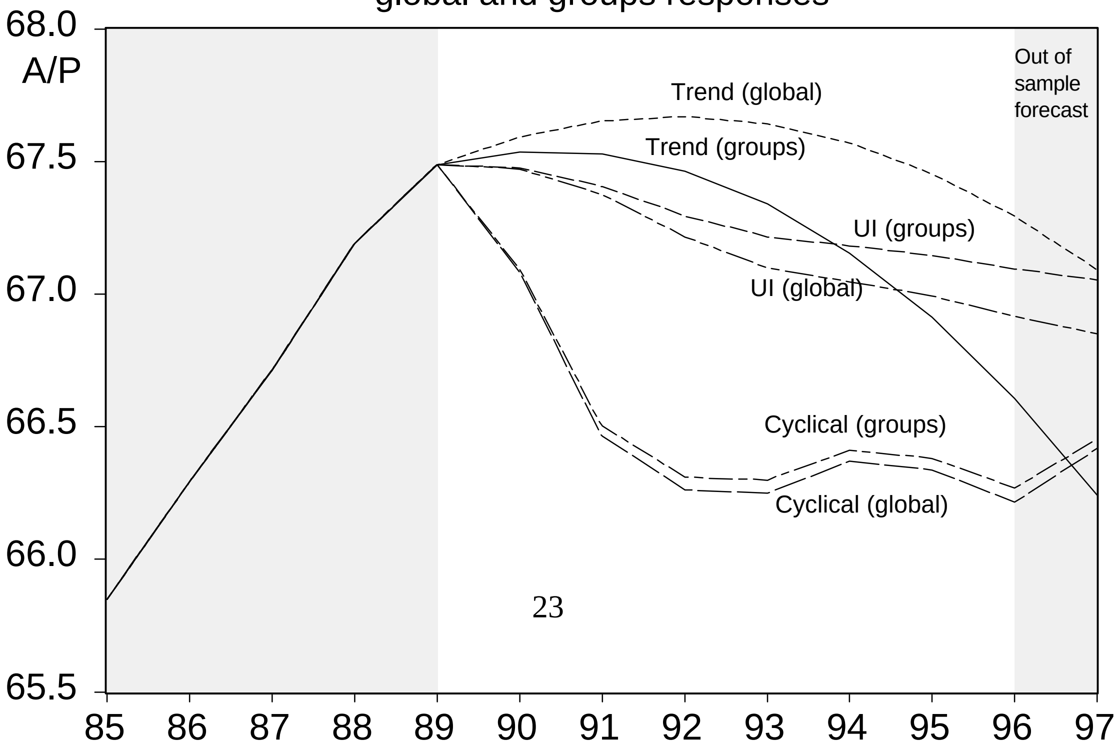

Figure 4
Decomposition
for males 15-24

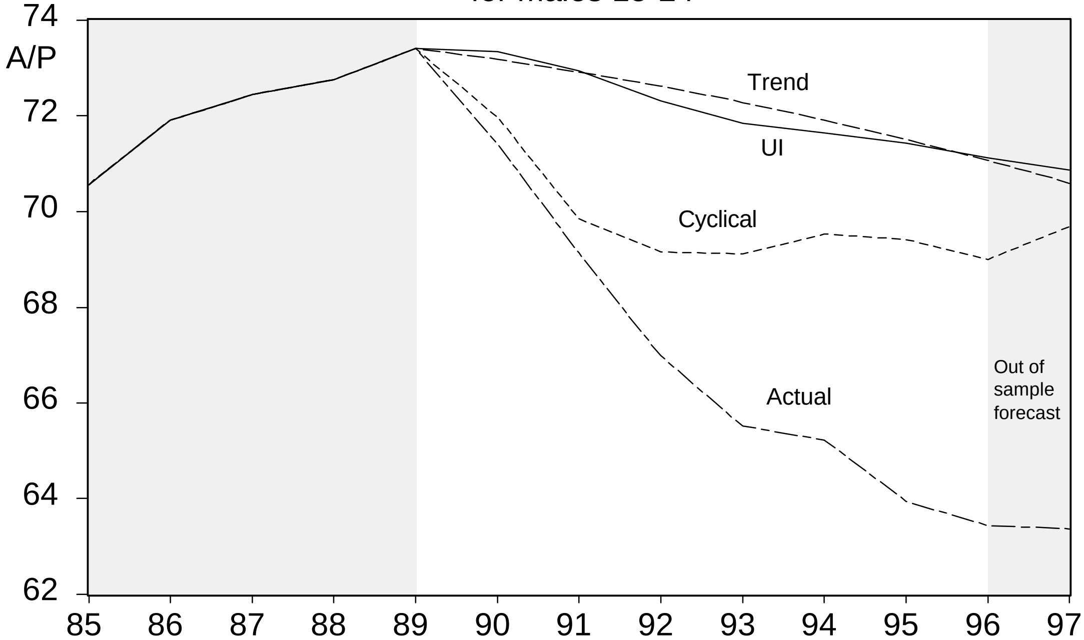

Figure 5
Decomposition for
females 15-24

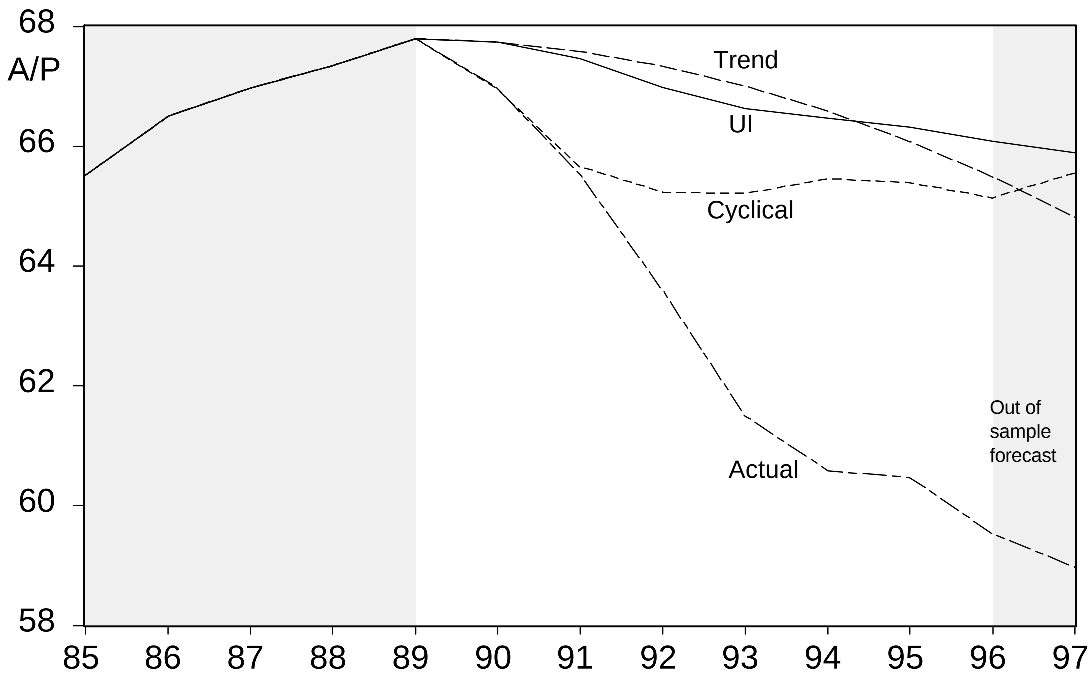

Figure 6
Decomposition
for males 25-54

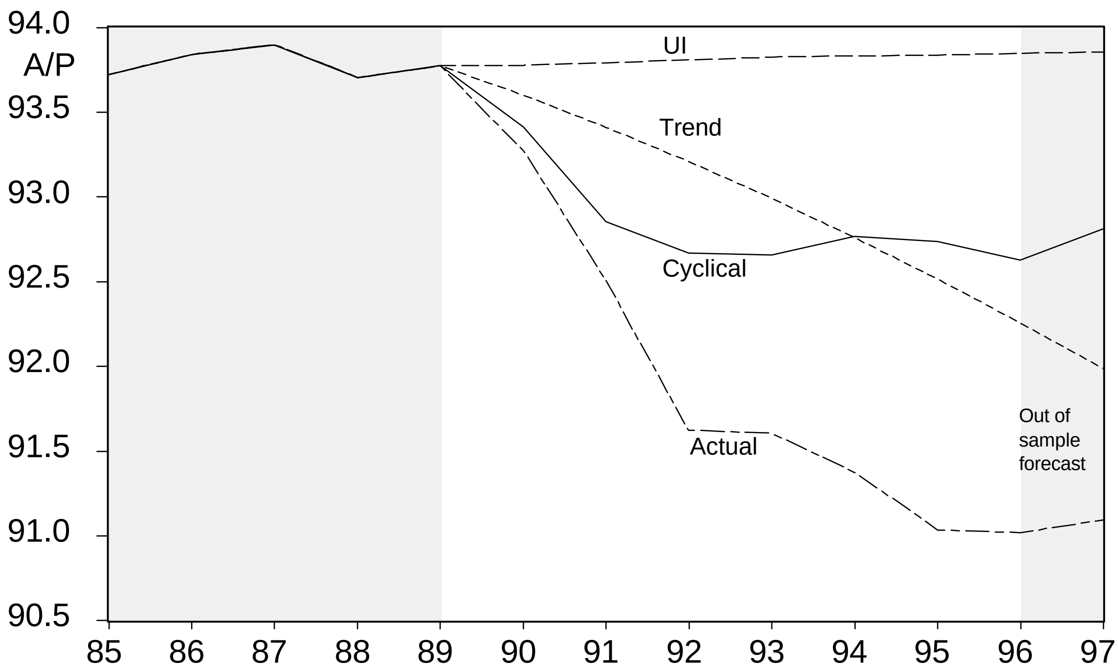

Figure 7
Decomposition
for females 25-54