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New Considerations on the Empirical Analysis of Health Expenditures in Canada: 1966–1998

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Résumé

L'objectif de ce rapport est d'analyser empiriquement les facteurs responsables de la hausse des dépenses de santé réelles par habitant au Canada et de déterminer si les dépenses de santé augmentent plus que proportionnellement par rapport à une augmentation du revenu, comme c'est le cas pour un bien de luxe (élasticité-revenu supérieure à l'unité).

Ces questions ont des implications de politiques majeures pour le Canada. Par exemple, s'il est établi que les soins de santé représentent un bien de luxe, les dépenses de santé en proportion du PIB augmenteront. Ceci veut dire que le secteur de la santé consommera une partie de plus en plus importante du revenu national. Puisque les revenus des gouvernements en proportion du revenu national peuvent difficilement augmenter, il s'en suit qu'une partie de plus en plus importante des revenus des gouvernements sera allouée aux dépenses de santé au détriment d'autres secteurs.

Afin d'étudier ces points, nous utilisons des données des 10 provinces canadiennes de 1966 à 1998 sur le revenu total réel par habitant, la proportion de la population âgée de 65 ans et plus, ainsi que le ratio du déficit ou surplus par rapport au PIB (produit intérieur brut). Des techniques économétriques récentes pour les données panel ont été appliquées afin de tester la stationnarité (si la moyenne, la variance et la covariance changent avec le temps) ainsi que la coïntégration (s'il existe une relation de long-terme entre les variables) des séries.

Nos résultats indiquent un coefficient d'élasticité de 0.88 pour le revenu, ce qui établit une relation positive entre les dépenses de santé et le revenu, mais ne suggère nullement que les soins de santé représentent un bien de luxe sous l'hypothèse des prix constants. Nous discutons également des implications du cas ou le prix relatif des soins de santé n'est pas constant. La recommandation de politiques qui en résulte est que plus d'argent doit être injecté dans le système de santé au fur et à mesure que notre économie se développe. Par ailleurs, un coefficient significatif de 1.45 a été trouvé pour le ratio du déficit ou surplus au PIB. Ceci suggère que, suite à une augmentation du déficit, les gouvernements peuvent être inclinés à réduire leurs dépenses dans les programmes pour diminuer le déficit. Nous avons obtenu un coefficient non significatif pour la proportion de la population âgée de 65 ans et plus, suggérant que l'effet du vieillissement sur la croissance des dépenses de santé a été jusqu'à date négligeable comparativement aux autres facteurs.

Classification JEL : I10, C22, C23, E30.
Mots clés : dépenses de santé, provinces canadiennes, stationnarité, coïntégration, données panel, élasticité revenu, effet Baumol.

Abstract

The purpose of this paper is to empirically study the factors that impact the Canadian provincial governments' real per capita health expenditures and to determine if health care expenditures in Canada grow more than proportionally with any increase in income, as in the case of a luxury good (income elasticity greater than one).

These issues have major policy implications for Canada. For example, if it is established that health care represents a luxury good, health expenditures as a proportion of GDP will increase. This means that the health sector will consume a larger and larger share of national income. Because it is hard for governments to increase their revenues as a proportion of national income, it follows that a growing share of governments' revenues will be allocated to health expenditures, at the expense of other sectors.

In order to address these issues, we use data on real per capita total income, the proportion of the population 65 years of age or over, and the ratio of the deficit or surplus to gross domestic product (GDP) for all 10 Canadian provinces for the period 1966 to 1998. Recent econometric techniques for panel data have been applied to test for stationarity (if the mean, the variance and the covariance of the series vary over time) as well as for cointegration (whether there is a long-term relationship between the variables) of the series.

An income elasticity of 0.88 is found, suggesting that health care expenditure is positively related to income, but that health is not a luxury good under the assumption of constant prices. We also discuss the implications of a situation where the relative price of health care is not constant. The resulting policy recommendation is that we need to put more money into the health care system as our economy grows. For the ratio of the fiscal deficit or surplus to GDP, a significant coefficient of 1.45 is obtained. This suggests that following an increase in the deficit, governments may be inclined to decrease their program expenditures in order to reduce deficit. The coefficient for the proportion of the population 65 years of age or over is non-significant, suggesting that the estimated effect of population ageing on health expenditure growth has been negligible relative to other factors to date.

JEL Classification: I10, C22, C23, E30.
Key words: Health expenditures, Canadian provinces, Stationarity, Cointegration, Panel data, Income elasticity, Baumol effect.

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1. Introduction

In the health economics literature, a debate has long raged over the determinants of increases in health expenditures. Using data from various OECD (Organisation for Economic Co-operation and Development) countries, different authors using various models have sought to determine the main factors explaining the increase in health care spending. The first generation of studies used cross-sectional regressions—initially bivariate (Newhouse, 1977) and later multivariate (Leu, 1986; Gerdtham et al., 1992). The second generation of studies consisted of panel data analyses. The earliest of these analyses (Gerdtham, 1992; Hitiris and Posnett, 1992; Gerdtham et al., 1998) did not consider the problems of non-stationarity (meaning series where the mean, the variance and the covariance vary over time) and cointegration that may exist between health expenditure and income variables. But the later ones did look at the issues of non-stationarity and cointegration between these variables (Hansen and King, 1996; Blomqvist and Carter, 1997; McCoskey and Selden, 1998; Roberts, 2000). One common finding in these studies is that real per capita gross domestic product (GDP) seems to be the most important factor in explaining the rise in health expenditures. But such studies have not yet managed to determine unanimously and conclusively whether health care services represent a "luxury good"—in other words, whether the income elasticity of these expenditures is greater than one. Some studies have found an elasticity coefficient greater than one, while other, more recent studies have found a coefficient of less than one. These contradictory results may be explained by the differences in the models. In addition to income, authors such as Leu (1986) and Gerdtham et al. (1998) have included other non-institutional variables such as the proportion of the population below age 15 or age 4, respectively; the proportion above age 65 or age 75, respectively; and alcohol and tobacco consumption. Except for tobacco, these non-institutional variables have been found to have little or no significance. Those authors have also included institutional variables, such as the number of physicians per 1000 population, the use of primary care "gatekeepers" and the degree of public sector involvement in the delivery of health care services. In most cases, results are presented with some caveats, Gerdtham and Jönsson (2000, pp. 45–47).

Like most of the major OECD countries, Canada has experienced a substantial increase in the proportion of GDP that it allocates to health expenditures. For example, for the period 1975 to 1998, the ratio of total health expenditures to GDP rose to 9.1% from 7.1% with a peak of 9.9% in 1992 (CIHI, 1999). The issue is thus just as crucial for Canada. The severe budget constraints of the Canadian provincial and federal governments during the first half of the 1990s further aggravated the problem of public funding for health services. In Canada, health expenditures are financed chiefly by the public sector[1] (jointly by the two levels of government). But under the provisions of the Canadian constitution, the provincial governments have the primary responsibility for providing health care services to their citizens.

1 In 1998, the Canadian average for provincial health expenditures as a proportion of total health expenditures was 64.4%. Some territories and provinces even had proportions as high as 75% that year.

This study will examine two questions:

- What factors are determining the rise in health care expenditures in Canada?[2]
- Does health care represent a luxury good (in the economic sense of the term) in Canada?

These questions have major policy implications for Canada and represent some of the most important issues in any study of the sustainability of the Canadian health care system. For example, if it is established that health care does represent a luxury good, health expenditures as a proportion of GDP will increase. This means that the health sector will consume a larger and larger share of national income. Because it is hard for governments to increase their revenues as a proportion of national income, it follows that a growing share of governments' revenues will be allocated to health expenditures, at the expense of other sectors. Despite all this, these issues have received very little attention in the empirical literature in Canada, except in Di Matteo and Di Matteo (1998), who attempted to use data for the Canadian provinces to determine the main factors explaining the rise in health care expenditures in Canada. The Di Matteo and Di Matteo study can be regarded as using a second-generation model: a panel data analysis in which the issues of group stationarity are not addressed.

The current study attempts to answer the two questions above using Statistics Canada data for the 10 Canadian provinces for 1966 to 1998, on provincial government health expenditures, total income (disposable personal income plus government income), the proportion of the population 65 years of age or over, and the ratio of the deficit or surplus to GDP.[3] It examines the effect of income on health expenditures in the context of relations between non-stationary panel series. This analysis has been facilitated by new tools for testing for the stationarity and cointegration of panel data (Im et al., 1997; Kao, 1999).

This study also takes a closer look at economic theory to separate conceptually two effects that are combined in the income-elasticity coefficient: the income effect and the Baumol effect. The latter is the idea that as the economy develops, certain labour-intensive services become relatively more expensive, while prices of manufactured goods fall. This effect appears to have been neglected in most empirical studies of health economics both in Canada and internationally. Health care is a labour-intensive sector. In Canada, approximately 60% of health expenditures are for labour (Carr and Ariste, 2001). The Baumol effect is therefore likely to come into play in this sector. That is why this study conceptually separates prices (which produce the Baumol effect) from quantities (which produce the income effect) in health expenditures.

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2 For a survey of some of the problems facing the Canadian health system, see Evans (1983); Culyer (1988); Brown (1991); Blomqvist and Brown (1994); Gratzer (1999).

3 Other variables such as the level of hospital use or drug use, or alcohol consumption could be included in the analysis, but time-series are not always available for these variables.

In summary, the current study differs from past ones in that it:

- focuses on Canada only and uses provincial data;
- includes fiscal variables to reflect government budget constraints;
- uses recently developed econometric methods to test the stationarity of panel data; and
- assigns a preponderant role to economic theory in explaining the relationship between health expenditures and income, and in particular for distinguishing income elasticity from the Baumol effect.

Nevertheless, we are aware that an estimated elasticity coefficient for Canada may be incorrect, because the price is not observed, and the health care demand function therefore cannot be calculated directly. That said, our results indicate a significant income elasticity coefficient of 0.88. This result suggests that at constant prices, health care represents a normal good and not a luxury good. However, the Baumol effect suggests that over time, there has been some variation in the relative prices of health care; this leads to a substitution effect that should be applied to correct the income effect obtained. As regards our two other variables, the coefficient found for the ratio of the deficit or surplus to GDP is equal to 1.45 and is highly significant, but the coefficient for the proportion of the population 65 years of age or over was not found to be significant.

The rest of this paper is organized as follows. In section 2, we present our data. In the first part of section 3, we present our model in which the time structure of the expenditure and income variables is formally considered. This leads us to test for the presence of a unit root,[4] first at the level of each province, then for all the provinces as a whole. We then perform tests for cointegration[5] of health expenditures and income for those provinces where these two variables were found to be integrated of order 1, or I(1).[6] In the second part of section 3, we try to determine whether an analysis based on the panel of observations (a pooled time-series cross-section approach) together with a good specification can help to obtain reliable results for the elasticity coefficient and the other parameters. Lastly, in section 4, we discuss the policy implications of our findings and present our conclusions.

4 The detection of a unit root indicates that we are dealing with non-stationary series. The application of the ordinary least squares method to such series would produce a spurious regression, indicating a false relationship.

5 If two series are non-stationary, there may be a long-term relationship that ties them to and prevents them from deviating very much from each other. In such a case, the two series are said to be cointegrated, and an error-correction model can be applied.

6 If a non-stationary variable becomes stationary after being differenced once, it is said to be integrated of order 1, or I(1).

2. Data

The variables considered in this study on the determinants of the health expenditures of Canadian provincial governments[7] are the total real income of the provinces, the proportion of the provincial population 65 years of age or over, and the ratio of the deficit or surplus to GDP. The reasons these variables were chosen are discussed below. The data series, originally stated in current dollars, was converted into real (1992) dollars using the Canadian GDP deflator. Because outcome measurements are not used in the Canadian health care system, deflators for health care products tend to be of very poor quality. The Canadian GDP deflator has been used in this study instead, because it expresses the opportunity cost of the other goods and services bought and sold in the country. Also, those series containing observations for a province as a whole were converted into observations per capita using the number of inhabitants of the province in question. We have chosen to use data on the scale of the provinces, and to pool these data into cross-section time-series (i.e. to use the pooled time-series cross-section approach), because it is the provinces that provide the health care services to their populations.

2.1 Health Expenditures of the Provincial Governments

The data used for health care expenditures by the provincial governments are those provided by Statistics Canada. We chose this source because it contains observations over a greater number of years. When looking for a cointegration relationship, a longer series is always better. Statistics Canada has been publishing annual data on health expenditures for each province since 1965/1966, whereas the Canadian Institute for Health Information (CIHI)[8] has such data only from 1975 on. Using Statistics Canada series on *consolidated revenue and expenditure of governments by function* (Statistics Canada historical data on public sector finance, Publication 68-512 and data for more recent years, Publication 68-212), we have obtained observations on health expenditures from 1966 to 1998. We thus approached more closely the conditions that let us make use of the theory of large numbers and the asymptotic properties of certain estimators. Figure 1 shows the average real per capita health expenditures for the 10 Canadian provincial governments. Real per capita health expenditures have increased from $351 in 1966 to $1,464 in 1998. Despite the decline between 1993 –and 1997, the average annual growth rate has been 4.5% between 1966 and 1998.

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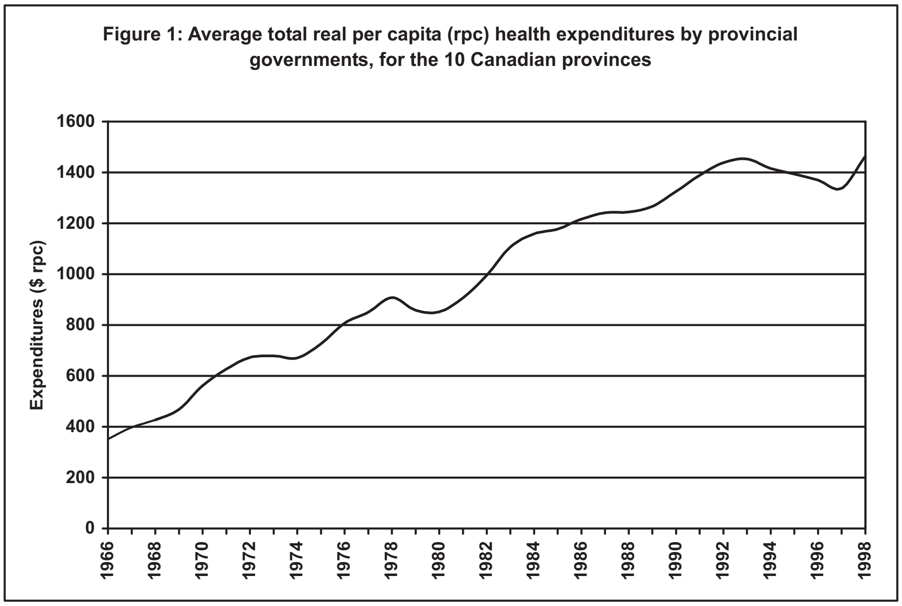
7 This study deals particularly with government expenditures on health care. To allow comparisons with international studies that use total health expenditures, results are provided later on using data on total health expenditures in Canada (see footnote 26).

8 It should be noted that Statistics Canada tends to apply a more restrictive definition of health expenditures than the CIHI. For example, Statistics Canada classifies health expenditures of the provincial ministries of social and community services as social services expenditures, whereas the CIHI includes them as health expenditures in its national health expenditures (NHEX) database. However, Statistics Canada and the CIHI do periodically reconcile their data on government health care expenditures.



2.2 Total Real per Capita Income

Real income is commonly included as a variable in empirical studies of the factors responsible for the rise in health expenditures. The income variable used in the present study has two components: disposable personal income and government income. Disposable personal income, which essentially represents GDP minus the taxes paid by individuals and businesses, is taken directly from the *Provincial Economic Accounts – Selected Economic Indicators*, CANSIM matrices 6968-6977 and 9220-9229. Government income comprises the taxes collected by the provincial governments plus the amounts of the transfer payments from the federal government to the provinces.[9] These data on provincial government income are taken from matrices 6769-6778 and 9085-9094. Adding these two income components, we obtain the total real per capita income for the provinces. In this way, we capture the provinces' spending ability more effectively than if we had considered GDP alone. Figure 2 shows average total real per capita income for the 10 Canadian provinces by the provincial governments.

As for health expenditures, this chart shows that real per capita income keeps growing with time. From $9,904 in 1966, it reached $21,634 in 1998 for an average annual growth rate of 2.4%**.**[10]

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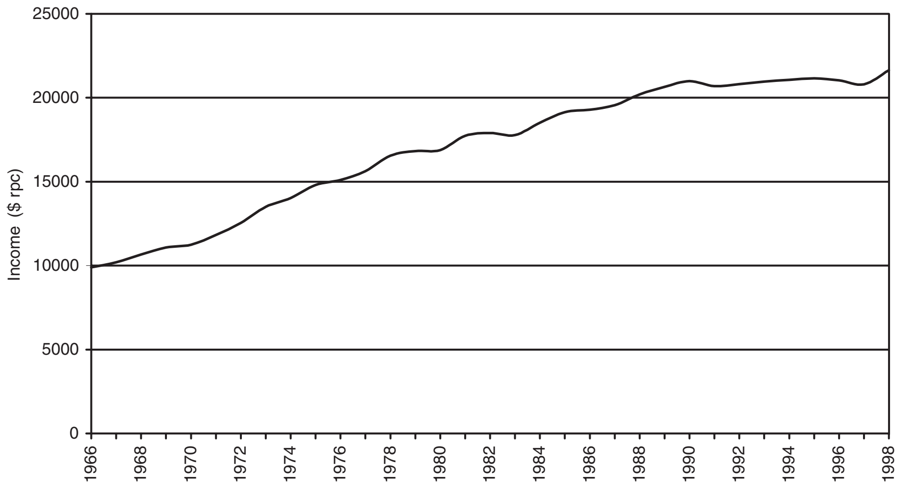
9 Di Matteo and Di Matteo (1998) consider the amounts of federal transfers to the provinces as an explanatory variable in and of itself. But we chose not to do so, because these amounts are not used exclusively to fund health services; in some cases, a large portion is predisposed via interest on public debt. We have treated them like any other source of income and added them to these other sources. One suggestion, with which we agree, is that it would have been appropriate to use government income and net out the portion of the transfers that go to debt servicing. We actually run a regression with provincial government income to which we net out provincial debt charges and provide the results in footnote 28.

10 Note that we cannot draw any conclusion from those charts and growth rates regarding the assumption of health as a luxury good. Only multiple regression can help to draw such conclusion.

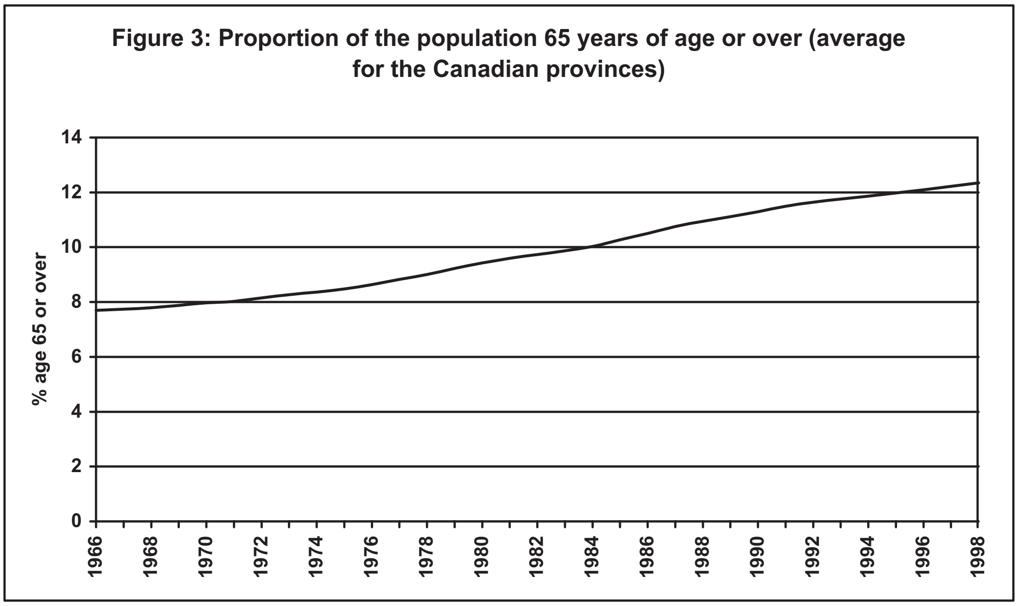
Figure 2: Average total real per capita (rpc) income by provincial governments, for the 10 Canadian provinces



2.3 The Proportion of the Population 65 Years of Age or Older

Our reasons for including the variable "proportion of the population 65 years of age or older" merit some explanation. Consumption of health care is not evenly distributed across the various stages of life; young children and seniors make more intensive use of it. In general, the costs run relatively high for very young children, drop considerably for people in their youth, then start increasing, at first gradually and later steeply, as people age. Denton and Spencer (1975) have shown that compared with health care expenditures for someone 40 years of age, those for someone 60 years of age are almost twice as high, and those for someone 70 years of age are almost three times as high. Likewise, Pollock (2000) has shown that the ageing of the population will contribute to a rise in health expenditures. The results of Pollock's projections indicate that the contribution of an ageing population to annual health expenditures will increase to about 0.9% around the time when the ratio of seniors to the total Canadian population peaks.[11] The ageing factor taken alone will thus contribute to increasing the proportion of GDP allocated for health expenditures. Even so, the effect will not be very great in terms of the contribution to nominal growth, and the study indicates that this factor will continue to play only a secondary role in the future. Figure 3 shows the evolution of the proportion of the seniors' population in Canada.

11 It is somewhat arbitrary to assign a date to the phenomenon of the baby boom. However, in Canada, the number of births peaked in 1959, and in subsequent years the birth rate fell rapidly. Consequently, it seems reasonable to say that all of the baby boomers will have reached retirement age by around 2030.



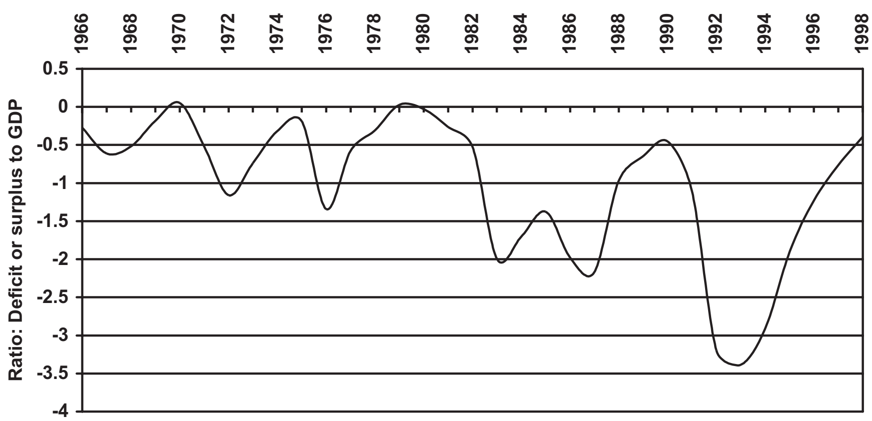
This graph shows that the proportion of seniors in the population is a trending variable also. Given that this proportion rose from 7.6% in 1966 to 12.35% in 1998 and, according to Statistics Canada's projections, should reach 21.4% in 2026, the large number of seniors can be expected to contribute to rising health expenditures.

The results of most empirical studies using international data show a non-significant coefficient for this variable (Gerdtham, 1992; Blomqvist and Carter, 1997; Barros, 1998; Roberts, 1998). In contrast, a study on Canada by Di Matteo and Di Matteo (1998) yielded a significant coefficient (but did not include the time trend found in those other studies). The current study, which uses the series *Population by single years of age, age groups and sex, by province* (CANSIM matrices 6368-6377) to obtain observations for this variable, will attempt to obtain another source of empirical estimates of the weight of seniors in the projection for health expenditures in Canada.

2.4 Ratio of Deficit or Surplus to Gross Domestic Product

Like Gerdtham and Jönsson (2000), we think that the ratio of the deficit or surplus to GDP could be an important element in determining health care expenditures. This variable has not been included in previous empirical studies on the determinants of health care expenditures. To capture the effect of governments' budgetary situation on health expenditures, we decided to include the ratio of the budget deficit or surplus to GDP in the regression. The sources that we used to calculate the data for this variable were the *Provincial government revenue and expenditure* (Public Institutions Division of Statistics Canada [SDDS 1720]) and *Gross domestic product, expenditure-based* (National Accounts and Environment Division of Statistics Canada [SDDS 1902/13-213[). Figure 4 outlines the trend in the average ratio of the budget deficit or surplus to GDP for Canada's provincial governments.

Figure 4: Ratio of budget deficit or surplus to GDP (average for the 10 Canadian provinces)



This graph shows that the ratio of the deficit or surplus to GDP evolves in a sawtooth fashion. After periods of budget deficits, governments tend to restrain their expenditures in an attempt to balance their budgets. The impact of such cutbacks on health expenditures can be tremendous, because they account for as much as 40% of some provincial governments' total program expenditures.[12] The reduction in health expenditures from 1993 to 1997 (see Fig. 1) goes along with a sharp reduction in the deficit for the same period.

12 In 1998, for Canada as a whole, provincial health expenditures averaged one third of the provinces' total program expenditures.

3. The Model

The basic model considered in this study is taken from the literature on the determinants of health expenditures. We postulate that real per capita health expenditures by provincial governments are a function of the provinces' real per capita income,[13] the proportion of the provincial population 65 years of age or over, and the ratio of the provincial budget deficit or surplus to GDP.[14] The model is specified in the form of a log-linear relationship between real per capita health expenditure and real per capita income. Thus we define:

- $h_{i,t}$ as the natural logarithm of the values of real per capita health expenditures for province i in year t;
- $r_{i,t}$ as the natural logarithm of the values of total real per capita income[15] for province i in year t;
- $a_{i,t}$ as the proportion of the population 65 years of age or over for province i in year t; and
- $b_{i,t}$ as the ratio of the provincial government's budget deficit or surplus to provincial GDP for province i in year t.

In this study, we attempt to account for the time structure of the expenditure and income variables. To do so, we must test for the unit root and, if all the series are non-stationary and I(1), for the cointegration of these panel data. The recent methods of Im, Pesaran and Shin (1997), hereafter IPS, and Kao (1999) are used to test for the unit root and for cointegration, respectively. The unit root tests are performed on each of the series using the augmented Dickey-Fuller method, hereafter ADF (Dickey and Fuller, 1981; Davidson and MacKinnon, 1993). We then obtain the statistic used to perform the unit root test for the panel by calculating the mean of the individual ADF statistics. This value is compared with simulated critical values provided by IPS. When this value exceeds a given significance level, the null hypothesis for the unit root is rejected.

The Kao method of testing for cointegration[16] consists of performing individual ordinary least squares (OLS) regressions of $h_{i,t}$ on $r_{i,t}$ and performing ADF tests on the estimated residuals for these series (Engle and Granger, 1987). The statistic used to test the null hypothesis of non-cointegration of the panel is obtained by calculating the mean of the ADF statistics

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13 One could think that the inverse relation is also true and that the simultaneity question should be addressed. However, it is unlikely that income is determined by the level of health expenditures.

14 Ideally, the deficit to GDP ratio should enter the regression with a lag, instead of contemporaneously. However, given our low sample size, loss of observations would even more hardly qualify our test statistics as asymptotically normal.

15 We also run a regression with provincial government income in which we subtract provincial debt charges. Results are provided in footnote 28.

16 This approach is not used, because we have shown later that one of the series is I(2).

previously obtained.[17] It is then compared with critical values supplied by Kao, and if it exceeds these values, the null hypothesis can be rejected. This therefore leads us to analyze the series for each of the provinces.

3.1 Analysis of Stationarity and Cointegration of the Series

The purpose of this section is to verify whether the variables in the regression are non-stationary and cointegrated. Because only $h_{i,t}$ and $r_{i,t}$ are *a priori* likely to be so, we consider only these two variables initially. If they are in fact found to be, we will then extend the analysis to the other variables in the model. We start first by testing whether each of the variables $h_{i,t}$ and $r_{i,t}$ is stationary. To do so, we use the ADF test which is based on the following regression for each series:

$$\Delta h_t = \alpha + \theta t + \beta h_{t-1} + \sum_{j=1}^{p} \gamma_j \Delta h_{t-j} + e_t \qquad (1)$$

$$\Delta r_t = \alpha + \theta t + \beta r_{t-1} + \sum_{j=1}^{p} \gamma_j \Delta r_{t-j} + e_t \qquad (2)$$

where e_t, for t =1,2,…,n is white noise; the number of lags p is chosen so as to eliminate autocorrelation of the residuals and minimize the Akaike Information Criterion. The regressions used to test the stationarity of the level variables can include a constant and a linear trend. Whether or not the linear trend should be included is the subject of growing debate.[18] For our part, we have performed the unit root test using an equation that incorporates both representations (i.e. a formulation with a constant [a stochastic representation] and a time trend [a deterministic representation]). The non-rejection of the null hypothesis for the unit root indicates that the series is characterized by a random walk representation. Also, only the constant is included in the regressions used to test the stationarity of the first-differenced variables. Table 1 shows the results of the unit root test on the values of $h_{i,t}$ and $r_{i,t}$ as well as their first differences. MacKinnon's critical values for testing the null hypothesis for the unit root at the 5% and 10% levels are -3.573 and -3.220, respectively, for the level variables, and -2.959 and -2.618 for the first-differenced variables. Ljung-Box Q statistics for testing the null hypothesis of the absence of order 9 autocorrelation in the residuals are also supplied, with the probabilities of rejection in parentheses.

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17 Note that this test allows you to have ordinates at the origin and different slopes between the groups.

18 McCoskey and Selden (1998) indicated that the ADF regressions should not include any linear trend, because the intercept itself already acts as a trend and power is lost in the case of a limited sample. These authors found that the health expenditures and GDP of the OECD countries are stationary. Hansen and King (1998) argued that the time trend is evident for these variables and must be included to apply the ADF test in its general form; they found the opposite result with the same data for the OECD countries. Also, Gerdtham and Löthgren (2000), using the KPSS test and the same data, rejected the null hypothesis of stationarity even when the trend is excluded.

At the 5% significance level, the null hypothesis of the unit root cannot be rejected ($\beta = 0$) for $h_{i,t}$ and $r_{i,t}$ in any of the provinces. However, at the 10% level, the null hypothesis of the unit root is rejected for $h_{i,t}$ in Ontario and Manitoba.[19] For the first differences of $h_{i,t}$, the null hypothesis of the unit root is rejected for each of the provinces, at both the 5% and the 10% levels; this suggests that the values of $h_{i,t}$ are I(1), because their first differences are stationary. As regards the first differences for $r_{i,t}$, they are stationary in seven of the ten provinces at the 5% level and in nine of the ten provinces at the 10% level; they are non-stationary in Alberta. In fact, the income variable for Alberta is I(2), because the second difference is stationary with an ADF statistic of -6.41. Moreover, according to the Ljung-Box statistics, the null hypothesis of non-autocorrelation cannot be rejected in any of the cases, which shows that the lags have been chosen appropriately.

Table 1: Unit root tests for health expenditures and income

Province	$h_{i,t}$	**Q-**$h_{i,t}$ **(Prob.)**	$<h_{i,t}$	$r_{i,t}$	**Q-**$r_{i,t}$ **(Prob.)**	$<r_{i,t}$
Newfoundland	-1.789	9.13 (0.425)	-5.678*	-1.045	15.75 (0.072)	-3.527*
Prince Edward Island	-2.237	12.96 (0.164)	-5.763*	-2.141	13.88 (0.126)	-4.318*
Nova Scotia	-2.933	5.35 (0.803)	-5.149*	-0.043	9.63 (0.383)	-4.319*
New Brunswick	-2.287	7.15 (0.62)	-3.759*	-0.38	7.83 (0.554)	-3.348*
Quebec	-2.374	6.74 (0.664)	-3.581*	-1.128	7.26 (0.612)	-2.931**
Ontario	-3.252**	4.34 (0.887)	-2.974*	-0.738	5.00 (0.835)	-2.925**
Manitoba	-3.487**	4.92 (0.841)	-4.729*	0.032	5.38 (0.803)	-5.158*
Saskatchewan	-0.928	3.52 (0.94)	-4.586*	-2.010	14.93 (0.09)	-3.853*
Alberta	-1.263	9.85 (0.362)	-3.455*	-1.423	6.99 (0.638)	-2.269
British Columbia	-2.172	12.13 (0.206)	-3.480*	-0.985	9.20 (0.420)	-3.324*

** , ** represent 5% and 10% significance levels, respectively.*

19 In a finite sample, the standard ADF test tends to reject the null hypothesis less often than it should have. The 10% level could help, to some extent, to offset this problem. Phillipes-Perron tests were also performed for purposes of comparison, and they provided the same results.

To verify the stationarity of the group and compensate for the weak power of the ADF tests when applied to small samples, we used the IPS method that proposed a unit root test in the context of a panel model using the mean of the individual ADF statistics for regressions (1) and (2). Our cross-section time-series data should ideally satisfy the hypotheses necessary to apply the alternative t-bar statistic[20] allowing us to test the null hypothesis of the unit root for all values of i ($\beta_i = 0$):

$$\bar{t}_{NT}(p_i) = \frac{1}{N}\sum_{i=1}^{N} t_{iT}(p_i) \qquad (3)$$

where $t_{iT}(p_i)$ represents the ADF tests estimated with p_i lagged differences, N is the number of groups, and T is the total number of observations.

The hypotheses are:

1. The number of groups (N) and the number of observations (T) are finite and satisfy the criterion $N/T \rightarrow K$ for $K > 0$;
2. The deterministic component is the same for all the groups; these groups have the same trend. For example, if we assume that the series h_t for one of the provinces has a linear trend, none of the provinces can have a stochastic component, which is, however, restrictive;
3. Every group considered in this study has a unit root, which renders plausible the alternative hypothesis of the IPS test that none of the groups has a unit root; and
4. The residuals from the individual ADF regressions are not autocorrelated, which is the case since the number of lags was chosen in such a way that the residuals would be white noise.

IPS propose using the following standardized statistic:

$$Z_i = \frac{\sqrt{N}\left(\bar{t}_{NT} - E\left(\bar{t}_{NT}\right)\right)}{\sqrt{Var\left(\bar{t}_{NT}\right)}} \qquad (4)$$

where $E(t\text{-}bar_{NT})$ and $Var(t\text{-}bar_{NT})$ are respectively the arithmetic means and the variances of the individual ADF statistics, given $\beta_i = 0$. They can be calculated using Monte Carlo simulations and are in fact supplied by IPS. The study by IPS shows that this statistic converges weakly toward the standard normal distribution for N tY, which lets it be compared with the critical values of the distribution N(0,1).

20 The other test proposed by IPS is based on the mean of the Lagrange multiplier statistics calculated for each group.The LM test of the unit root has been considered by Solo (1984) for time series.

In our case, because N is not large enough, we cannot use this Z statistic. Instead, we have used the IPS critical values in a finite sample. These values are -2.58 and -2.49, respectively, for the 5% and 10% levels of the values of N and T in the study. The values calculated for the t-bar statistical tests are -2.268 for the set $h_{i,t}$ and -0.893 for the set $r_{i,t}$. On this basis, we conclude that each of the variables $h_{i,t}$ and $r_{i,t}$ forms a non-stationary group.

The next step in our analysis is to check whether all the pairs $h_{i,t}$ and $r_{i,t}$ are cointegrated for each province, except for Alberta[21]—in other words, whether there is a long-term relationship between health expenditures and income. One way to proceed is to use the method of Engle and Granger (1987)[22] which consists in using the OLS to estimate the equation:

$$h_t = \alpha_0 + \beta_0 r_t + u_t \qquad (5)$$

and to subject the residuals from this estimate to unit root tests. The rejection of the null hypothesis of the unit root constitutes evidence of cointegration.

Table 2 shows the results of the estimation and the ADF tests on the residuals. These tests are done with a constant but with no trend. MacKinnon's critical values for the rejection of H_0 at 5% and 10% are -2.959 and -2.618, respectively. The statistical evidence indicates that at the 5% significance level, the null hypothesis of non-cointegration is rejected in all cases except in Newfoundland, Quebec and Saskatchewan. At the 10% level, the null hypothesis is rejected in all cases. If we consider the 5% level, we may suggest that health expenditures and incomes are non-stationary and cointegrated in three of the four Atlantic Provinces, as well as in Ontario, Manitoba and British Columbia. Given the weakness of the ADF tests with small samples, we propose to find another means of cross checking these cointegration relationships. One of the sufficient conditions for a true cointegration relationship is that the residuals of regression (5) be non-autocorrelated.

21 It should be borne in mind that income in Alberta is I(2).

22 We could also use the method of Johansen, which is more robust when there are several cointegration vectors. This method tests the hypotheses on these cointegration vectors directly.

Table 2: Tests for cointegration between health expenditures and income

Province	α_0	β_0	ADF(u_t)	R^2o-AC1 (Prob.)
Newfoundland	-6.01	1.33	-2.909	9.83* (0.007)
Prince Edward Island	-6.60	1.38	-4.403*	2.64 (0.104)
Nova Scotia	-8.98	1.64	-3.636*	5.94* (0.015)
New Brunswick	-7.701	1.50	-3.648*	6.63* (0.010)
Quebec	-8.92	1.62	-2.753	16.831* (0.000)
Ontario	-13.27	2.05	-3.648*	16.59* (0.000)
Manitoba	-10.97	1.83	-5.191*	10.45* (0.001)
Saskatchewan	-5.66	1.28	-2.675	17.76* (0.000)
British Columbia	-14.18	2.14	-3.64*	12.62* (0.000)

** 5% significance level*

The last column in Table 2 presents the results of the Breusch-Godfrey test of the Lagrange multiplier to test the null hypothesis of the absence of first-order autocorrelation (AC1), with the probabilities of rejection shown in parentheses. The observed R^2 statistics suggest that at the 5% level, the test is significant in all cases, except for Prince Edward Island, which indicates that H_0 is rejected and the residuals are autocorrelated in most of the cases.[23] The fact that we have cointegration and autocorrelated residuals simultaneously in certain provinces represents an anomaly. It may be due to the small size of the sample, because the ADF tests are powerful only asymptotically. One way of getting around this problem would be to group the provinces and use the method of Kao (1999) which tests the null hypothesis of non-cointegration of panel data. This is not possible, however, because the income series are not I(1) for all of the provinces.[24] We therefore conclude that the panel data on health expenditures and income in Canadian provinces are non-stationary and possibly non-cointegrated. In the context of individual series,

23 This is the proof that the model is poorly specified and that we are dealing with a spurious regression in this case.

24 If we discount the case of I(2) in Alberta and test for the cointegration of the panel, we find a t-bar of -3.575. According to Kao's critical values, we will reject the null hypothesis of non-cointegration at 5% and 10%, but not at 1%. However, these results are not considered in this study.

we could explore a VAR model[25] but we prefer to pool the data in order to meet the conditions that let us make use of the asymptotic properties of the estimators. This leads us to consider a cross-section time-series model taking into account not only a linear trend, but also the autocorrelation and the groupwise heteroskedasticity of the errors.

3.2 Results for the Panel Model

Due to the weak sample size, we use a panel model to take advantage of all the information available in the data and to have more degrees of freedom and power. The model we use includes a trend and corrects for autocorrelation and groupwise heteroskedasticity of the errors. We also introduce the variables $a_{i,t}$ and $b_{i,t}$ (Ref, p. 18) to obtain the following complete model:

$$h_{i,t} = \alpha_i + \theta t + \beta r_{i,t} + \gamma a_{i,t} + \eta b_{i,t} + e_{i,t} \qquad (6)$$

where α_i is regarded as a term that is constant over time and specific to the group (a fixed effect) and the values $e_{i,t}$ are the residuals of the idiosyncratic or characteristic effects. This model thus imposes the following four restrictions:

$$\theta_i = \theta;\ \beta_i = \beta;\ \gamma_i = \gamma;\ \eta_i = \eta \qquad (7)$$

We estimated this model using the Stata 7 *xtgls* command, which is appropriate for estimating cross-section time-series models and simultaneously allows the inclusion of the *panels (heteroskedastic)* options to specify the heteroskedastic structure with no cross-correlation of the errors and *corr(ar1)* to specify that within the panels, there is first-order autocorrelation.[26] The results are reproduced in Table 3.

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25 The VAR model could be more compatible with the stochastic representation found in the series but, taking into account the relatively few numbers of observations, it would create some concern in using the theory of large numbers.

26 We first estimated this model with the hypotheses of homoskedasticity and non-autocorrelation of the errors. Not surprisingly, these hypotheses were rejected.

Table 3: Results for panel model with linear trend

$h_{i,t}$	Coefficient	Std. Dev.	z score	P > |z|
$r_{i,t}$	0.88	0.069	12.671	0.000
$a_{i,t}$	-1.35	0.726	-1.868	0.062
$b_{i,t}$	1.45	0.204	7.121	0.000
t	0.022	0.002	9.180	0.000
constant	-2.00	0.642	-3.113	0.002

Because the health expenditures and income variables are expressed as natural logarithms, the coefficient of $r_{i,t}$ represents elasticity. This coefficient suggests that, all other things being equal, a 10% increase in income should result in an 8.8% increase in health expenditures. When we test the null hypothesis that $r_{i,t} > 1$, we reject it at the 5% significance level with a chi-square (1) of 2.94.[27] This indicates that health represents a normal good and not a luxury good[28]: that is, if we assume that preferences and prices remain constant, the demand for health care is positively related to income, and the increase in the demand for health care is less than proportional to any increase in income.

However, it is far from certain that the "price" of health care is constant. The 8.8% increase in health expenditures is the combined result of increases in the two components of health expenditures: quantities (of equipment, supplies, labour, etc), which produces the income effect, and prices (of equipment, supplies, labour, etc), which produces the Baumol effect. To capture the income effect alone, we would need historical data on the quantities of all the inputs used in the health care system, and we would have to regress these quantities onto income and the other variables. Or else, to separate the Baumol effect from the combined effect, we would need data on the wages of health professionals, and we would have to regress them onto income and the other variables. This would be beyond the scope of the present study.

However, according to Carr and Ariste (2001), the increase in labour expenditures in the health sector is due not to the increase in the quantity of labour, but mainly to the rise in wages per unit time and per worker.[29] This suggests that the assumption of constant prices does not hold in the case of health care. Because the health sector is labour intensive, it experiences wage increases that result in non-negligible cost increases. Given the inherent limitations on increasing production in the health sector, the possibilities of productivity gains are therefore limited. This

27 In fact, the null hypothesis tested is $r_{i,t} = 1$. It is rejected for any significance level exceeding 8.64%.

28 Since provincial governments are largely responsible for health expenditure in Canada, we also run the regression with government income net of debt charges. The coefficient on $r_{i,t}$ drops to 0.34, suggesting health is far from being a luxury good. The coefficients for the other variables $a_{i,t}$, $b_{i,t}$ and t are –1.21, 1.42 and 0.033, respectively.

29 This is confirmed in the case of labour but not necessarily for the other inputs.

situation results in an increase in the relative prices of the labour of medical staff. While the sign of the variation in relative prices is known for labour, it is not known for the other inputs of health care, such as drugs or technology.[30]

The increase in the demand for health care following an increase in income will be lesser (8) or greater than it would have been if prices had remained constant; whether lesser or greater depends on the sign of the variation in the relative prices of health care. Assuming labour costs have a determining effect on output prices, the relative prices are rising (Baumol effect). If we let Δx be the change in the quantity of units of health care demanded and ΔP_x be the change in the price of a unit of health care, we can write:

$$\frac{\Delta x}{\Delta r} \left| \Delta P_x > 0 \right. \quad < \quad \frac{\Delta x}{\Delta r} \left| \Delta P_x = 0 \right. \qquad (8)$$

Baumol effect: underestimation of the income effect

$$\frac{\Delta x}{\Delta r} \left| \Delta P_x < 0 \right. \quad > \quad \frac{\Delta x}{\Delta r} \left| \Delta P_x = 0 \right. \qquad (8')$$

Improvement in technology or increase in productivity: overestimation of the income effect

The result can be either an underestimate (8) or an overestimate of the income effect obtained in regression (6). Discriminating between these two effects is important for the correct interpretation of the income effect. The issue is only raised in this study.[31]

One unexpected result of this regression is that the coefficient for the proportion of the population 65 years of age or older is negative, though not significant at the 5% level.[32] While contrary to that obtained by Di Matteo and Di Matteo, who found a positive, highly significant coefficient, these authors did not include the time trend unlike others (Gerdtham, 1992; Blomqvist and Carter, 1997; Barros, 1998; Roberts, 1998) who got a result similar to ours. If we rerun our own regression equation but exclude the linear trend, we too can obtain a positive, highly significant coefficient of 2.32. However, because the linear trend is excluded, this coefficient may also be capturing the effect of other variables that are not considered in the model. As for the unit root tests where the inclusion or omission of a trend can lead to different results (see footnote 16), the presence or absence of the linear trend is critical in determining the relationship between the ageing of the population and health expenditures.

30 Hence, in the final analysis, we do not know the sign of the variation in relative prices for all health care inputs as a whole.

31 A thorough discussion of the Baumol effect requires discussion about the technology of production in the health sector, something that is well beyond the scope of this paper.

32 We also estimated this model using the CIHI data on total health expenditures (1970–1998) and including the trend, and we were able to obtain a coefficient that was positive though also non-significant for age (even at the 10% level). The other variables $r_{i,t}$, $b_{i,t}$ and t are all significant with values of 0.56, -0.36 and 0.018, respectively.

The variable $b_{i,t}$ is highly significant, with the usual sign[33]: an increase of 1 percentage point in the ratio of the deficit to GDP corresponds, all other things being equal, to an increase of 1.45% in health expenditures. The impact of the 1990s has probably intensified this result (see Figure 4). Following an increase in the deficit, governments may be inclined to decrease their program expenditures in order to reduce deficit. The lagged budgetary balance could be a best indicator of the health expenditures level.

The coefficient of the linear trend suggests that even if income remains unchanged, health care expenditures in the Canadian provinces increase by up to 2% per year. This result may be interpreted partially as the contribution of the impact of technological change on growth in health expenditures.[34] More and more health economists believe that this factor may represent the most important determinant of health care expenditures.[35]

Regression (6) also provides the Wald statistic that can be used to test the null hypothesis that the four restrictions stated in (7) are valid. The chi-square of 2024 obtained constitutes strong evidence for rejecting this hypothesis, which makes us doubt the validity of the constrained model and the common coefficients obtained. An estimate by individual province might prove to be the best approach. Because of the small size of the sample, the results of the unit root and cointegration tests would have to be investigated in depth before the adoption of any model that applied this approach.

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33 Because the variable “ratio of deficit or surplus to GDP” takes negative values, the negative of this variable should be considered for a standard interpretation of the results.

34 Other factors that are probably masked in this coefficient are the changes in the relative prices of health care and in the health of the population.

35 For an overview of the role of increasingly sophisticated medical technologies in rising health expenditures, see Weisbrod (1991) and Feeny (1994).

4. Implications and Conclusion

In this study, we have examined the effects on health expenditures of three variables: income, the proportion of the population 65 years of age or older, and the ratio of the deficit or surplus to GDP. After analyzing the questions of stationarity and cointegration of the data series, we opted for a cross-section time-series model that takes into account the linear trend, the heteroskedasticity, and the autocorrelation of the error terms. Our results clearly indicate that health expenditures in Canada will continue to rise as the economy develops and we become wealthier; this is consistent with spending behaviour on a normal good. However, due to the uncertainty of the sign of the variation in the relative prices of health care, we cannot conclude whether or not health care services represent a luxury good, because the income coefficient obtained may underestimate or overestimate the "true income effect."

To establish whether or not this is so, it would be necessary to be able to empirically dissociate growth in unit prices of health care (the Baumol effect) from growth in the number of units of health care (the income effect). With regard to human resources, this would mean separating growth in wages per worker from growth in the number of workers. Lastly, the rejection of the constrained model suggests that an estimate of the individual elasticities from a vector-autoregressive model (VAR) or an error-correction model (ECM) could give us a better picture of the situation in the future when more data are available.

The health expenditures of the provincial governments will grow by 2% per year even without income growth. This may be due in part to the availability of medical technologies that are increasingly sophisticated and increasingly expensive to acquire. Hence, there is the potential for a budget deficit whenever growth in GDP is weak or non-existent; otherwise, the Canadian health system could have trouble remaining at the cutting edge of technology. Besides, the deficit is positively correlated with health expenditures: the value of the coefficient $b_{i,t}$ suggests that a 1 percentage point increase in the deficit-to-GDP ratio will be accompanied by a 1.45% increase in health expenditures.

All of these results are consistent with what might be expected intuitively. When the economy is growing, wages increase substantially. The impact on health expenditures can be enormous, because this is a labour intensive sector. As the primary providers of health care, the provincial governments acquire the best equipment available to ensure that they can provide their citizens with high quality services. But in acquiring this equipment, they do take their budgetary constraints into account, which may explain the observed lag in the technology of the Canadian health system compared with that of the United States.

The negligible impact of the ageing of the population on health expenditures represents a less conventional result for a Canadian study. Due to the difficulty of separating the ageing factor from that of the time trend, we cannot draw any firm conclusions for the moment. Whatever the case, this coefficient cannot be used to project the effects of ageing on health expenditures since the historical data did not capture the impact of the baby boomers on health expenditures. It seems likely that the pressure of retired baby boomers on the health system will be considerable. But it will not be as intense as many studies on this subject suggest, especially if we consider the nominal data.

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