The Determinants of Canadian Provincial Health Expenditures: Evidence from Dynamic Panel

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ABSTRACT

This thesis aims to reveal the magnitude of the income elasticity of health expenditure and the impact of non-income determinants of health expenditures in the Canadian Provinces. Health can be seen as a luxury good if the income elasticity exceeds unity and as a necessity good if the income elasticity is below unity. The motivation behind the analysis of the determinants of health spending is to identify the forces that drive the persistent increase in health expenditures in Canada and to explain the disparities in provincial health expenditures, thereby to prescribe sustainable macroeconomic policies regarding health spending. Panel data on real per capita GDP, relative price of health care, the share of publicly funded health expenditure, the share of senior population and life expectancy at birth have been used to investigate the determinants of Canadian real per capita provincial total, private and government health expenditures for the period 1975-2002. Dynamic models of health expenditure are analyzed via Generalized Instrumental Variables and Generalized Method of Moments techniques. Evidence confirms that health is far from being a luxury for Canada and government health expenditures are constrained by the relative prices. Results also cast doubt upon the power of quantitative analysis in explaining the increasing health expenditures.

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CHAPTER 1

INTRODUCTION

This thesis aims to reveal the magnitude of the income elasticity of health expenditure and the impact of non-income determinants of health expenditures in the Canadian provinces. Health can be seen as a *luxury good* if the responsiveness is sensitive to income changes (i.e. the income elasticity exceeds unity) and as a *necessity good* if the responsiveness is insensitive to income changes (i.e. the income elasticity is below unity). This concept was used by Newhouse (1977)[1]. Another interpretation of this notion can be found in Kyriopoulos and Souliotis (2002):

> *"If the income elasticity of health expenditure is less than one, then the public health sector does not have a high priority among the goals for social and economic development."*

The motivation behind the analysis of the determinants of health spending is to identify the forces that drive the persistent increase in health expenditures in Canada and to explain the disparities in provincial health expenditures, thereby to prescribe sustainable macroeconomic policies regarding health spending. Further, identifying the effects of income and institutional factors on public and private health expenditures allows inference about the trends in the public-private mix in Canadian health sector. The structure of this mix has been the centre of the debate of whether increasing centralization or privatization would yield more efficient outcomes[2]. Although inferences are being made regarding the trends in the public-private mix of Canadian health expenditures, assessing the efficiency of single or two-tier health systems is beyond the scope of the thesis.

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[1] The term "luxury" coined by Newhouse (1977) may sound as if health is "bad" rather than "good". This term is used to indicate that the elasticity is greater than one without referring to health being literally a "luxury".

[2] Di Matteo (2000) provides an excellent discussion on the public-private mix of Canadian health expenditures.

Estimating the impact of income and other measures on health spending forms the basis of this analysis. If in fact the income elasticity of health spending is less than one, then health expenditures would increase at a lower rate than GDP and the public health sector must not have high priority among the goals of economic and social development. Income elasticity below unity further implies that health is a necessity good and thus the delivery of health is determined according to needs[3]. To address these issues, this thesis focuses on the demand side of health care and on the determinants that are quantitative in nature rather than factors that measure the quality of life and health. Conclusions drawn from a demand approach aim to answer the following questions:

- Is health care in Canada a necessity or a luxury good?
- How important are the demand measures in explaining provincial health expenditures?
- Can demand measures be effective in order to control health spending?

Issues that are not considered in this analysis consist of production, supply factors, cost efficiency and medical technology among others.

1.1 Canadian Regional Health Expenditures at a Glance

The real per capita total and private health expenditures are shown in Figure 1.1 and 1.2[4]. Real per capita total health expenditures have shown an increasing trend until 1992, reaching an average of 2,400 dollars. Between 1975 and 1996 the average growth rate of the Canadian per capita total health spending was around 2.8%. The increase in health spending halted between 1992 and 1996. The Canadian average[5] growth rate in this period was slightly below zero. The flattening of real per capita total health expenditures in this period was caused by the economic downturn and the

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[3] See Di Matteo (2003).

[4] In this section, the graphs are analyzed regionally for the exposition but the data is analyzed at the provincial level.

[5] Section 1.1 commonly refers to the Canadian averages but to keep things simple they are not shown on the graphs.

increasing public deficit resulting in cuts in health expenditures[6]. After 1996, total health spending soared with an average growth of 4.4% and rose above 3,000 dollars by 2002. The Atlantic Provinces have the lowest and British Columbia has the highest per capita spending on health.

Figure 1.1: Real per capita total health expenditures
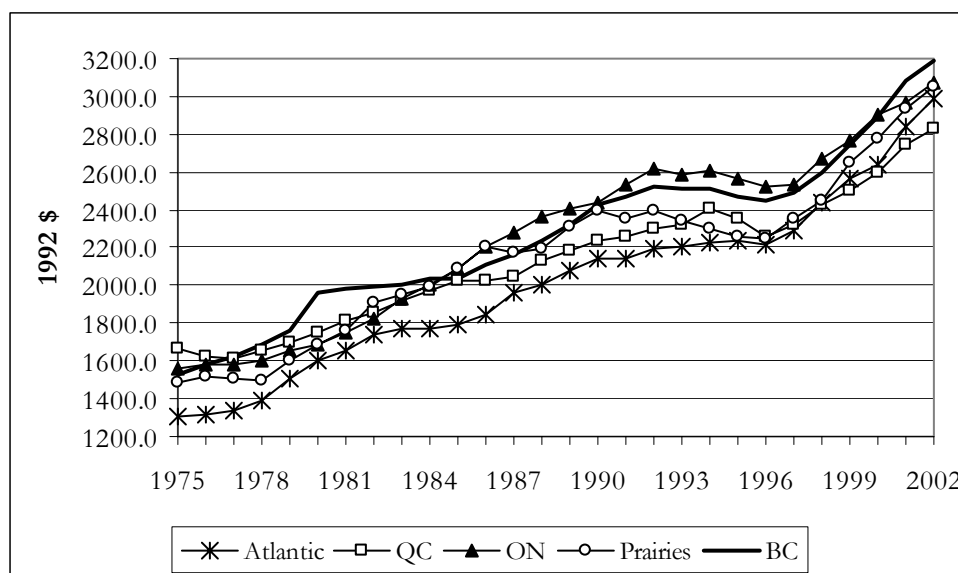
Until mid 1990s, there were slight variations in per capita private health spending across provinces. British Columbia and Ontario have had per capita private health spending above those of the rest of the provinces. However, in 1991 the provincial private health spending clustered with the exception of Ontario where the private health spending kept its steady rise thereafter. By 1997, the amount of per capita private health spending equalized across the Atlantic Provinces, the Prairies,

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[6] See Organization for Economic Co-operation and Development (OECD) 2001 report, "Policy Brief: OECD Health at a Glance - How Canada Compares"

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Quebec and British Columbia, reaching 650 dollars on average. Ontario is the province with the highest per capita private health spending of over 1,000 dollars.

Figure 1.2: Real per capita private health expenditures

Figure 1.3 and 1.4 displays the total and private health expenditures as a percentage of GDP. Quebec, Ontario, Prairies and British Columbia devote a share to total health expenditures below the Canadian average whereas the Atlantic Provinces devote a larger share of provincial GDP in both figures. For private health expenditures in the Atlantic Provinces, this share suddenly rose in 1976 and peaked at 3.7% by 1980 (see Figure 4). The steady growth of real per capita private drug expenditure in the Atlantic Provinces appears to be one of the causes of rising share of private health expenditures out of GDP until 1980. Between 1976 and 1980, the growth rate of real per capita private drug expenditure in the Atlantic Provinces was

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8.7%. But this growth of private drug expenditure was only 2.2% between 1981 and 2001.

Figure 1.3: Total Health Expenditure as a Percentage of GDP (%)

Source: Derived by dividing provincial total health expenditures (Canadian Institute for Health Information, table B.1.1) by provincial GDP (Statistics Canada, table 384-0001).

Figure 1.4: Private Health Expenditure as a Percentage of GDP (%)
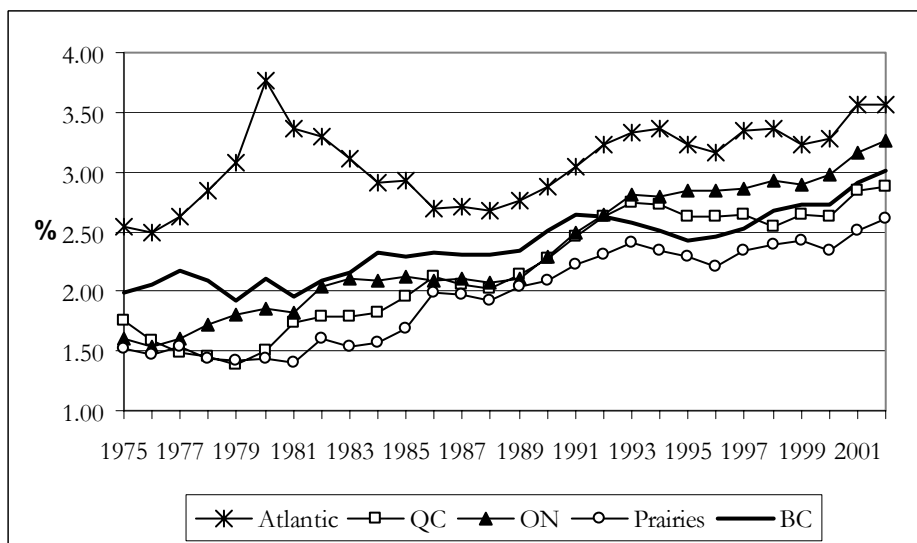
Source: Derived by dividing provincial private health expenditures (Canadian Institute for Health Information, table B.2.1) by provincial GDP (Statistics Canada, table 384-0001).

From 1979 to 1980, total health expenditures as a percentage of GDP increased by more than 1% and concurrently the share of publicly funded health expenditures in the Atlantic Provinces reached its lowest (see Figure 1.5). The share of private health expenditures out of GDP in the Atlantic Provinces dropped to 2.7% by 1986.

The share of publicly funded health expenditures[7] has always been above the Canadian average for Quebec and the Prairies. There also has been a downward trend in this share in Quebec and the Prairies since late 1970's, early 1980's and in Ontario after 1992. By 2001, the highest share of publicly funded health expenditure belongs to Newfoundland (80.5%) and the lowest share belongs to Ontario (66.1%).

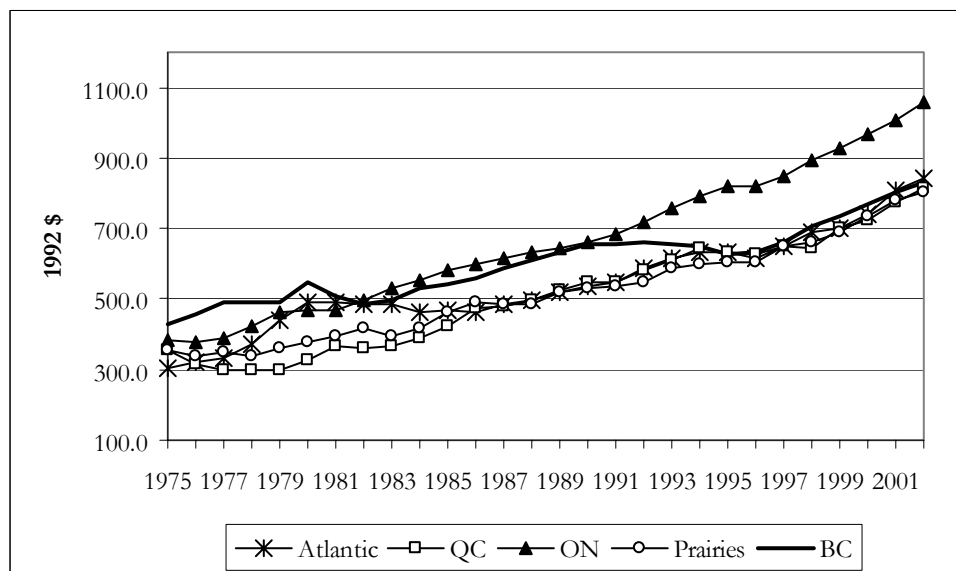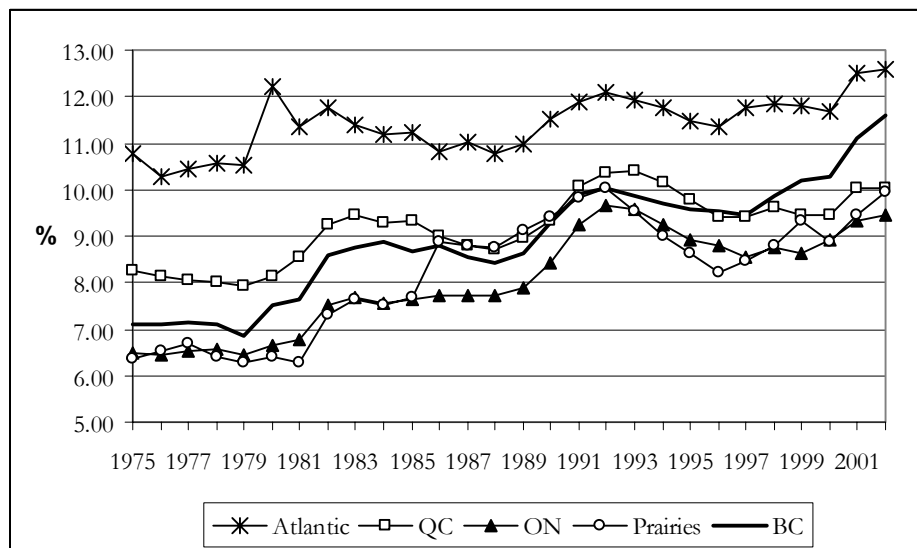**Figure 1.5:** Publicly Funded Health Expenditures as a Percentage of Total Health Expenditure (%)
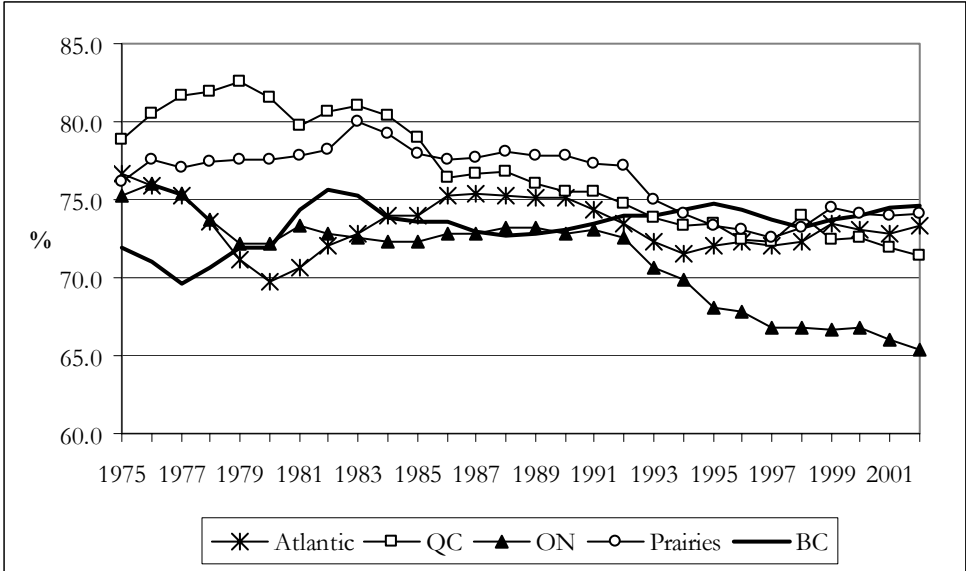
Source: Derived by dividing provincial public health expenditures by provincial total health expenditures (Canadian Institute for Health Information, tables B.1.1 and B.3.1).

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[7] Public health expenditures consist of expenditures by government and government agencies. A detailed content of public sector health expenditures can be found at Canadian Institute for Health Information website: http://secure.cihi.ca/cihiweb/dispPage.jsp?cw_page=statistics_nhex_definitions_e

1.2 Literature Review

The analysis of the determinants of health expenditures (HE) has been very tempting for both applied econometricians and health economists. Nevertheless, there is no consensus on which method(s) to use, how to proceed and what type of data to analyze. This may have occurred due to lack of strong theoretical guidance. The pioneering studies emphasize the importance of national income in explaining HE along with a selection of non-income variables. Some of these variables are the relative price of health care (ratio of medical CPI to GDP price index), the proportion of the population over the age of 65, urbanization rate and the publicly funded proportion of HE among others. While the significance of non-income variables depends on the structure of health sector and population, GDP accounts for most of the variation in aggregate health care expenditure – see Parkin et al. (1987).

Newhouse (1977) asked the question "what determines the quantity of resources a country devotes to medical care?" His study led to a body of literature on the determinants of health care expenditure. His regression of per capita medical care expenditure on per capita income for 13 United Nations countries yielded income elasticity greater than one, suggesting that health is a luxury good. The results also suggested that the estimated income elasticities decrease with rising income level.

Some of the studies in the early 1990s considered cross-section analysis. Gbesemete and Gerdtham (1992) used cross-sectional data for 30 African countries to measure the effects of socioeconomic and demographic variables where per capita HE is a function of percentage of births attended by health staff, per capita GDP, percentage of the population under 15 years of age, urbanization rate, crude birth rates and per capita foreign aid. The results showed that only per capita GDP, percentage of births and per capita foreign aid were statistically significant and carried a positive sign. In two out of three estimated equations the coefficients of GDP were below unity. Gerdtham et al. (1992) investigated a similar relationship for 19 OECD (Organisation for Economic Co-operation and Development) countries using cross-section data and generalized a model where HE is a function of national income,

relative price of health care, institutional factors, age structure and urbanization. The results indicated that the income elasticity is greater than one. Gerdtham et al. (1994) incorporated socioeconomic and demographic factors such as GDP, age structure, alcohol and tobacco consumption and female labor force participation ratio as well as various institutional factors, depending on the structure of health sector in 22 OECD countries. For all of the four models, they found that income elasticity is significantly below unity.

There is a growing panel literature on the determinants of HE in the OECD countries[8]. Those studies relied on pooled cross-section and time-series data of OECD member countries to partially overcome the small-sample shortcomings. Gerdtham (1992) used pooled cross-section and time-series data for 22 OECD countries and compared different models of HE including a static equilibrium model, ECM (error correction model) and dynamic models such as ARDL (autoregressive distributed lag), growth rate and partial adjustment. The results indicated that the short-run income elasticity is below unity whereas the long-run income elasticity of HE is around unity. Hitiris and Posnett (1992) analyzed the determinants of HE using a sample of 560 panel observations for 20 OECD countries. While the results support that GDP is the most important determinant, they arrived at the conclusion that the income elasticity is larger than one. They also concluded that non-income variables are important but their effects on HE are small. Moore et al. (1992) specified a model for cross-country examination where per capita HE is a function of per capita income, per capita number of physicians, nurses and beds, and the ratio of public expenditures to total health care expenditures. The results indicated that the number of per capita beds has a negative effect on health care spending. They also found that health is a necessity in the short-run while a luxury in the long-run.

Hansen and King (1996) employed country-by-country analysis and postulated that the HE-GDP relationship is spurious if the variables are not stationary[9]. They also

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[8] See Roberts (1999), Okunade and Karakus (2001), Di Matteo (2003) for example.
[9] See Section 3 for a discussion on stationarity and unit roots.

argued that a non-linear specification should not be ignored by just relying on log-linear functional forms of HE models. Kanavos and Mossialos (1996) emphasized that the inclusion of national income does not reflect the society's ability to pay. They argued that GDP may not be effective at all in explaining HE (their study confirmed that Germany, France, Ireland, Netherlands and Greece provided evidence of insignificance of GDP in national currency). Blomqvist and Carter (1997) used panel data for 18 OECD countries for the period 1960-1991 to check the common finding that health is a luxury good. They found out that both real per capita total health expenditure and income are non-stationary and cointegrated for most of the countries. The long-run income elasticity estimated separately for each country turned out to be around or greater than unity for all countries with the exception of Canada, UK and the USA. Pooling of OECD countries with country specific intercepts yielded an income elasticity close to unity. Casasnovas and Saez (1998) examined the factors involved in rising health care expenditures by developing a model using data for 110 regions in eight OECD countries in 1997. They identified two sources of variation (within countries and between countries) to find out if the relationships between health care spending and the explanatory variables are country-specific. The model is specified as a mixed fixed and random coefficients model (MFR) by allowing the constant and the income parameter to differ randomly across countries where log of per capita health expenditure depends on log of per capita income, share of public health spending and the share of population over the age of 65. The results confirmed that the income elasticity is far less than one. They also found that increases in the share of public spending and share of population over the age of 65 are associated with increases in health spending. Hitiris (1999) examined the forces that drive the rising health expenditures in the G7 countries and the cost containment policies. His findings confirmed that income and the share of senior population explain almost 90 percent of the variation in health spending.

As longer time-series of HE became available for the OECD countries, the dynamic properties of the relationship were captured in a *data field* framework.

Roberts (1999) concentrated on the shortcomings of the analysis of health care expenditure and employed techniques to analyze dynamic heterogeneous *data field* with non-stationary variables for 20 OECD countries over the period 1960-1993. She specified an ARDL random coefficients model to capture the heterogeneity across the OECD countries where total health spending is a function of income, proportion of the population over the age of 65, relative price of health care, the proportion of publicly funded health spending and a time trend which captures technological change. She compared mean group, pooled and cross-section estimators. In static mean group and pooled estimation she found evidence of significant long-run effects of income, the proportion of publicly funded health spending and the relative price of health care. Only the long-run income elasticity was significant in the dynamic mean group estimation. She also focused on sensitivity analysis to check the robustness of the results and the parameter sensitivity to country exclusion. The reported long-run income elasticity was above unity. In a later study Roberts (2000) addressed the problem of spurious regression in Hitiris' 1997 article. She warned that the regression of non-stationary variables results in misleading correlation arising from the common trends in the data rather than the true economic relationship. Robert's cointegration approach using Hitiris (1997) dataset yielded a short-run income elasticity significantly below unity whereas Hitiris found an income elasticity above unity. However, the existence of any long-run relationship between health spending and its determinants was unclear. Getzen (2000) argued that health care is an individual necessity and a national luxury in the sense that the magnitude of income elasticity depends on the level of analysis and researchers fail to distinguish between sources of variation *within groups* and sources of variation *between groups*. At the individual level, budget constraints do not provide sufficient information about how much to spent on health care as long as there is a system of pooling resources which removes those individual constraints. However, national constraints still exist and Getzen argued that the analysis should be based on the units of observation at which decisions are being made.

Okunade and Karakus (2001) employed individual ADF and IPS panel unit root tests, Engle-Granger and Johansen cointegration analysis for real per capita health expenditure, real per capita GDP and relative price of health care in 19 OECD countries between 1960-1997. They estimated a GDP elasticity of HE above one, yielding consistency with the previous estimates. A case study for the UK revealed that health is a luxury good with income elasticity of 1.43 and the responsiveness of the UK health spending to changes in the relative prices is found to be highly elastic. Bac and Le Pen (2002) focused on estimating a demand function by using panel data for 18 OECD countries by adopting a cointegration approach where per capita health expenditure depends on per capita GDP and the relative price of health care. They have found strong evidence on the cointegration of these variables and compared OLS, fully modified OLS (FMOLS) and dynamic OLS (DOLS) where the latter two estimators account for endogeneity and serial correlation. The results confirmed that the income elasticity of health care spending exceeds unity. Clemente et al. (2002) examined the stability of HE models in the OECD countries by adopting a cointegration approach. They criticized the stability assumption of HE-GDP relationship and argued that there exist structural breaks which lead to a biased and incorrect long-run relationship[10]. They conducted the analysis by disaggregating total expenditure as public and private health expenditure. The results suggested that the inclusion of structural breaks does not invalidate income elasticity of health care spending being greater than one.

There exist other empirical studies in the OECD countries examining the cointegrating relationships and unit root problem exclusively. McCoskey and Selden (1998) employed country-by-country ADF tests and IPS panel data unit root tests for per capita HCE and GDP in 20 OECD countries and they rejected the null hypothesis of unit root. They also concluded that one need not be concerned about the existence of unit root in the OECD data. Gerdtham and Löthgren (2002) used panel data for 25

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[10] Structural break refers to jumps in the series due to sudden policy changes, wars, institutional or macroeconomic shocks (see Maddala and Kim, 1998).

OECD countries for the period 1960-1997 and presented new results on cointegration of HE and GDP. While the unit roots tests indicated that both HE and GDP are difference stationary series, they found that in 12 countries out of 25, HE and GDP were cointegrated. Jewell et al. (2003) presented evidence on the stationarity and the presence of structural breaks in HE and GDP covering 38 years for 20 OECD countries. The panel unit root tests with heterogeneous structural breaks suggested on the contrary that both HE and GDP are stationary if they allow for structural break(s).

Di Matteo and Di Matteo (1998, henceforth DD) focused on the determinants of Canadian provincial government health expenditures within pooled time-series cross-section framework for the period 1965-1991. The determinants of provincial government health expenditures are found to be the real per capita provincial income, the share of senior population and real provincial per capita federal transfers. Although the issue of stationarity is not fully addressed, they reported that the income elasticity of government health care spending is 0.77. Di Matteo (2000) examined public and private Canadian health expenditures over the period 1975-1996. The major determinants of public-private mix are per capita income, the share of individual income held by the top quintile of the income distribution and federal health transfers. Health expenditures are examined as total and sub-expenditure categories such as hospital, physician and drug spending. The empirical evidence suggested that increases in per capita income are associated with more private health care spending relative to public spending. In a later study Di Matteo (2003) compared parametric and nonparametric estimation methods for the U.S states, the Canadian provinces and the OECD countries. He argued that parametric approach leads to unreliable estimates of the income elasticity of health expenditure as it assumes a particular functional form and its magnitude is highly dependent on the level of analysis. International level analyses lead to estimates greater than one in parametric approaches. Di Matteo also provided evidence that the income elasticity of health spending is higher at low-income levels and lower at high-income levels. Ariste and Carr (2001, henceforth AC) used provincial data on real per capita income, the proportion of the population over

the age of 65 and the ratio of the deficit/surplus to GDP to explain the real per capita government health expenditures. They examined the non-stationarity of the variables, the cointegrating relationships and found that variables, both individually and collectively, are non-stationary and possibly non-cointegrated. The determinants of government health expenditures are found to be income, the ratio of the deficit/surplus to GDP, the share of senior population and a time trend capturing technological progress. However the coefficient of the share of senior population appeared to be statistically insignificant. After concluding that all the variables are non-stationary and possibly non-cointegrated, they estimated a fixed effects model with non-stationary variables and concluded that the income elasticity of government health spending is 0.88. Table 1 summarizes the estimates of the income elasticity of HE in both regional and international contexts with the basic features of the analyses.

CHAPTER 2

DATA AND ECONOMETRIC METHODOLOGY

Few points that are not considered by DD and AC are of interest. First, if the relative price of health care is known to have an influence on HE (see Bac and Le Pen, 2000 and Okunade and Karakus, 2001 for example), the failure to incorporate this variable as one of the determinants will ultimately lead to specification bias and incorrect estimates due to the combined income and price effects. It should be noted that the income coefficient due to the exclusion of the health price variable may be biased either direction[11]. Second, the previous studies on the determinants of Canadian provincial health expenditures can be characterized by lack of dynamics. Income may have permanent and transitory components and the increments on income may not be fully spent in the same period but rather its spending may be allocated through time. Further, current period health spending may also depend on its past values. As Roberts (1999) mentioned, the structure of the adjustment process of health spending is not well known. Getzen (2000) argues that one should expect lags on the right-hand side as the budget is prepared at least a year in advance. These shortcomings indicate that the early estimates of the determinants of Canadian health expenditures may have been biased and conclusions drawn could have been misleading. It is argued that the dynamics of health expenditures should not be neglected for the purposes of modeling and policy implications. Third, the determinants of provincial health expenditures are examined in three categories: Total health expenditures, government health expenditures and private health expenditures. Disaggregation allows examination of the differing responsiveness against the income and price changes for the government and private sector as well as the total health spending. Fourth, other factors that are not

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[11] The direction of this bias depends on the sign and the magnitude of income and price elasticities.

considered in previous Canadian studies are incorporated in this analysis, such as the life expectancy at birth and the share of publicly funded health expenditures.

Finally, concerning unit roots, Atkins and Sidhu (2002) warned that if the series under consideration are not weakly stationary (i.e. the series contains at least one unit root) which is the case in most regional and international comparisons, then traditional econometric analysis is not valid. Failure to achieve weak stationarity will cast doubt on the statistical significance of the coefficients and their reliability. Even if economic theory weakly provides guidance on the relationship between HE and its various determinants, statistical theory shows that the mean and the variance of the underlying series should be time invariant (see Maddala and Kim, 1998).

2.1 Data Description

The data covers 10 provinces in Canada for the time period 1975-2002. However, only the observations between 1979 and 1999 are used in order to have a balanced panel[12]. The provincial total, the provincial private and the provincial government health expenditures are taken from the Canadian Institute for Health Information (CIHI) website[13]. These variables are deflated by the provincial CPI (1992=100) and divided by the provincial population to obtain real per capita provincial total (h), real per capita provincial private (pr) and real per capita provincial government (g) health expenditures. The share of publicly funded health expenditure (s) is obtained by dividing public health expenditures by total health expenditures. The provincial medical CPI (1992=100), provincial proportion of the population over the age of 65 ($p65$), life expectancy at birth (x) and the provincial GDP are collected from CANSIM[14]. The provincial GDP is deflated by the provincial CPI (1992=100) and divided by the provincial population to obtain the real provincial per capita GDP (y).

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[12] All variables used in this analysis cover the period of 1975-2002 with the exception of life expectancy at birth and the relative price of health care which cover the period of 1979-1999 and 1979-2002 respectively.

[13] http://secure.cihi.ca/cihiweb/dispPage.jsp?cw_page=statistics_results_source_nhex_e

[14] CANSIM table numbers are 384-0003, 102-0025, 326-0002, 051-0001, 384-0001.

The provincial medical CPI is divided by the provincial GDP price index (1992=100) to obtain the relative price of health care (*r*) for each province.

A couple of points need to be discussed in order to explain why the panel framework has been chosen. The first reason is the gain in sample size due to pooling of cross-section and time-series observations. This relatively large sample contains more information about the relationship between health expenditures and its determinants, allows investigation of dynamics and the disparities in provincial health spending, introduces more efficiency and more degrees of freedom. However, pooled estimation brings some issues. Dynamic pooled estimation, as will be shown in subsequent sections, requires some type of instrumental variable technique. In international comparisons, the pooled cross-section and time-series raised questions regarding the validity of homogeneity of health care demand functions, convertibility in unit of measurement and data comparability. These problems are less pronounced in a provincial comparison because of a unified health system across provinces and consistent data collection.

The outline of this thesis is as follows: Section 2.2 addresses the issue of non-stationarity with province-by-province and panel unit root tests. Section 2.3 introduces the dynamic health expenditure models and investigates the reasoning behind the relationship between health spending and its selected determinants. Chapter 3 discusses the results and the relevant policy implications. Chapter 4 summarizes the findings and concludes with directions for future research.

2.2 Province-by-province and Panel Unit Root Tests

Unit root is a severe problem in the sense that if the appropriate tests are not employed, the inferences drawn might possibly be misleading and "seemingly good" results may occur because of a common trend rather than true economic relationship (see Granger and Newbold, 1974). A model should be treated and interpreted over stationary forms of the variables. A common problem in time series is the existence of unit root. Most economic time series are classified as being integrated of order *d*,

denoted as *I(d)*, that is the series must be differenced *d* times in order to become stationary. Otherwise, regression of non-stationary variables results in spurious correlation. This study will first consider Augmented Dickey-Fuller (ADF) unit root test proposed by Dickey and Fuller (1979) under the null of unit root with its extension to panel by Im et al. (2003, henceforth IPS) and KPSS test proposed by Kwiatkowski et al. (1992) under the null of stationarity with its extension to panel data by Hadri (2000). See the appendix for technical discussion on individual and panel unit root tests.

The ADF and IPS results are shown in Table 2. The ADF results show that for most of the series of health expenditures, GDP and the share of publicly funded health expenditures, the null hypothesis of unit root cannot be rejected. Concerning total health expenditures, the null can only be rejected for New Brunswick, Prince Edward and British Columbia. In the case of GDP, this null can only be rejected for Prince Edward and British Columbia. The IPS panel t_{bar}-statistics show that all of the variables can be described as group stationary. It should be emphasized that concerning the IPS test, lag order or lag criteria greatly affects the individual unit root statistics in favor of rejecting the null hypothesis of unit root.

The KPSS individual unit root tests show that for most of the series except the share of senior population, the null of trend stationarity cannot be rejected. However, Hadri's panel unit root tests show that the null hypothesis of either level or trend stationary can be rejected for all the series at the 5% significance level. This result might be induced from the fact that the test proposed by Hadri is valid under sequential limit in which $T \to \infty$ followed by $N \to \infty$. The results are displayed in Table 3.

The first problem that appears in unit root testing is whether to include a time trend. While Hansen and King (1998) claimed that ADF regression should include a linear trend, McKoskey and Selden (1998) argued that it should not. This paper argues that most macroeconomic variables have tendency to increase over time, therefore it is appropriate to include a deterministic component into unit root testing. However, some

variables may not evolve around a trend component at all, yet may appear stationary. Economic theory does not help as to whether include a linear trend or not. At this point, we should rely on the statistical significance of the linear trend in unit root tests. The inclusion of such deterministic components is more or less heuristic.

Karlsson and Löthgren (2000) warned that unit root test such as IPS has high power in panels with large T and researchers might mistakenly conclude that the whole panel is stationary even though most of individual series are nonstationary. The converse is true if T is small. This argument is reconciled for both unit root tests that are undertaken. The decision concerning unit roots is inconclusive. For the IPS test, a significant fraction of the series is individually nonstationary but they appear to be stationary as panel. However, for Hadri's test a significant fraction of the series is individually stationary but they all appear to be nonstationary as panel. A careful assessment of individual and panel unit root tests should be undertaken to identify the order of integration of the variables with confidence. However, this is beyond the scope of this paper. It should be underlined that the presence of structural breaks is not considered by the unit root tests due to short time span of the series.

Our primary concern is whether or not the relationship between the Canadian HE and its determinants would be spurious if one analyzes this relationship in levels of the variables. From an economic point of view, shocks to the Canadian health sector have temporary effects rather than effects that alter the level of expenditure permanently. Thus, this analysis will proceed by assuming that the panel is weakly stationary and that the regression is unlikely to be spurious in level. Further, even if this is not the true case, any indication of spurious regression can be captured by the estimation results.

2.3 Dynamic Models of Health Expenditure
2.3.1 Factors Affecting Health Expenditure

Before introducing the models, this section discusses the reasons behind the inclusion of the selected factors into the analysis of provincial health expenditures.

The studies on the determinants of health expenditures argued that **income** is the major explanatory factor of HE[15]. The economic approach argues that *other things being equal*, the amount of health expenditure should depend on what an individual is capable of spending. Therefore it is expected that provinces with higher income should spend more on health taking other factors as given[16]. Figure 2.1 displays the provincial real per capita GDP. Differences in the provincial GDP should be taken into account when analyzing health expenditures as the Atlantic Provinces have much lower income levels compared to the rest of Canada. Alberta is the wealthiest province with real per capita income around 33,000 dollars on average followed by Ontario and British Columbia.

Figure 2.1: Real per capita GDP ($)
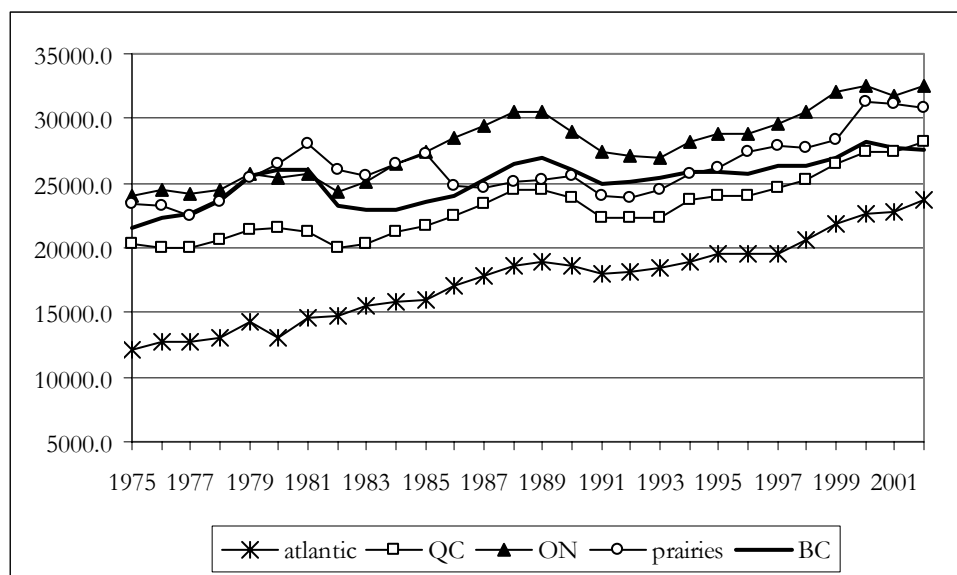
Source: Derived by dividing provincial GDP by total population and deflated by the CPI (1992=100). Statistics Canada tables 384-0001, 051-0001 and 326-0002.

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[15] See Newhouse (1977) and Parkin et al. (1987) for example.

[16] If individuals spend more on health as income rises, the aggregate expenditure on health consists of sum of *N* such individuals. However, it is argued that macro and micro outcomes are generally different due to insurance or pooling of resources or other factors in which case micro constraints may exist but macro constraints may not. These micro and macro disparities may render different or inconsistent outcomes. See Getzen (2000) for a discussion on individual versus aggregate.

Spending decisions concerning health are not solely affected by the income level but also by **the price of health care**. Especially in the case of higher out-of-pocket payments, decisions rely on the price level. On one hand the government is heavily involved into the delivery of health and its supervision. On the other hand, health care has special characteristics that are different than those of other *"goods"*. Such features pose problems about our expectations of the magnitude of the price effect and its sign[17]. This variable is particularly included into the analysis to separate income and price effects. From the economic point of view, the failure to include the price variable, if effective, results in misleading inference regarding policy prescriptions. The relative price of health care in the Prairies has been the lowest across Canada, increased between 1981 and 1993 with an average growth rate of 1.7% in that period. Figure 2.2 clearly shows that the medical CPI was growing faster than the overall price level until early 1990s. This was expected considering the burden of the cost of medical technology and the labor intensity in health care sector.

Figure 2.2: Relative Price of Health Care
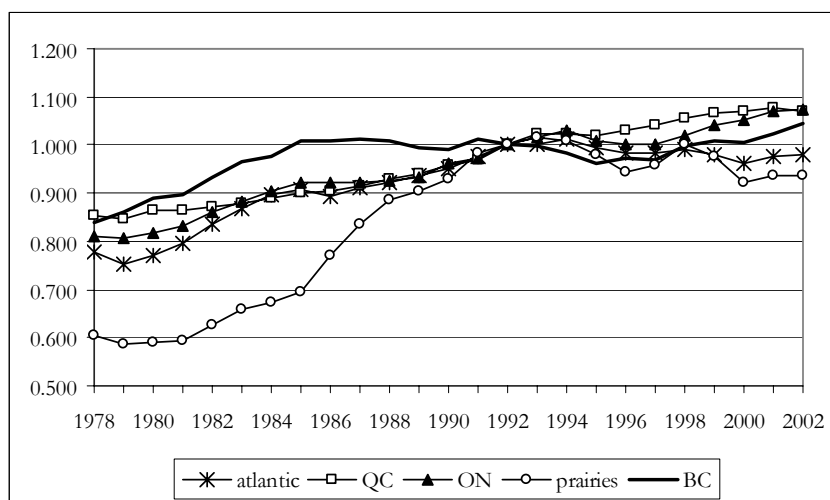
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[17] The first counter-argument to its inclusion is that the consumers never face prices for the health services they receive and therefore this variable may be completely irrelevant for the analysis. Secondly, price of health is heavily subsidized in Canada so that even its effect is not zero, it should be almost zero or negligible.

With few countries as exceptions[18], health care decisions and a considerable volume of health spending are driven by the governments and public institutions. Therefore, we expect **the share of publicly funded health expenditure** to affect health spending. If this share is effective in explaining the private heath spending, then inferences can be made regarding the interaction between public and private spending. However, as Roberts (1999) pointed out, both theory[19] and empirical evidence are contradictory regarding the magnitude and the sign of this effect.

The share of senior population is considered to be another explanatory factor of HE. The elderly population consumes health at a higher rate than others and the depreciation rate of health is an increasing function of age (see Grossman, 1972). Especially for those over the age of 65, higher and prolonged periods of cost are involved. The treatment of senior's population involves complexity and the elderly patients are not completely cured in most of the cases. Diabetes, cardiovascular diseases are just a few examples that require relatively technical knowledge and equipment for treatment and diagnosis. The delivery of health services to an elderly population is therefore associated with higher spending on health. However, a recent study at the individual level by Felder et al. (2000) presented evidence that people tend

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[18] The obvious example is the United States where the government has a smaller role in health sector in comparison to Canada.

[19] If *T, G* and *P* denote the real total, public and private health expenditures respectively and since $T = G+P$ the share of publicly funded health expenditures can be defined as:

$$\psi = \frac{G}{G+P}$$

Then the following equality holds:

$$\frac{\partial T}{\partial \psi} = \frac{\partial G}{\partial \psi} + \frac{\partial P}{\partial \psi}$$

The effect of a marginal increase in ψ on private health expenditure is:

$$\frac{\partial P}{\partial \psi} = \frac{P(G+P)}{G} = \frac{PT}{G} > 0$$

The effect of a marginal increase in ψ on total health expenditure is:

$$\frac{\partial T}{\partial \psi} = \frac{(G+P)^2}{G} > 0$$

The theory shows that the sign of this effect is always positive and the higher the private HE the greater the effect of ψ. Therefore, if the empirical evidence is consistent with the theory we would expect this effect to be positive. See Casasnovas and Saez (1998), Hitiris (1999) and Roberts (1999) for empirical evidence on the magnitude of this effect.

to spend more on health in their last 2 years of lives before death and within this period health expenditures decrease with age for individuals over the age of 65.

Figure 2.3: The Share of Senior Population (%)
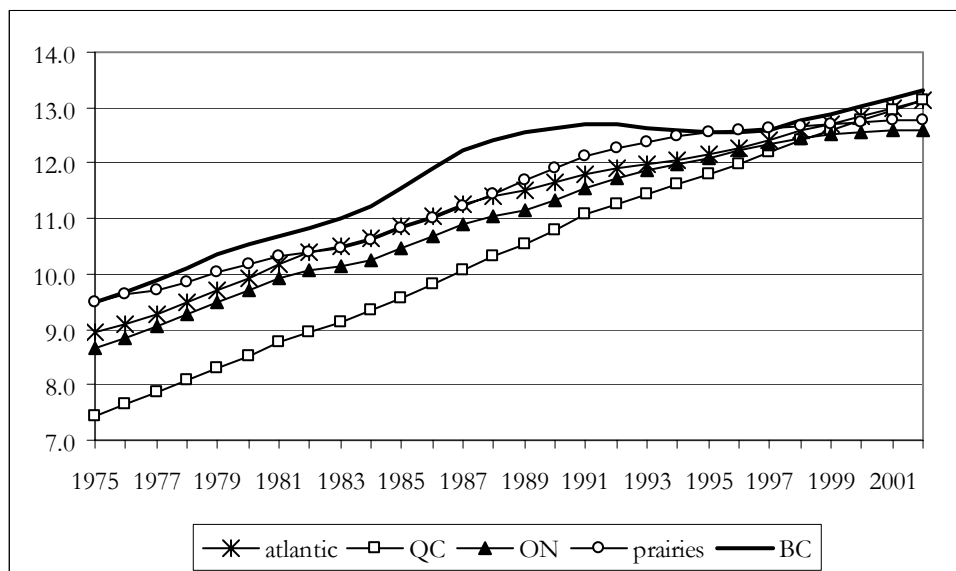
From 1975 until the end of 1980's, the average growth rate of the share of senior population across Canada was around 2%. But this growth has been slowed down, even turned negative in British Columbia towards the beginning of 1990's. However, the Canadian population consists of a growing share of the elderly, representing 12-13% of the population in 2002.

The relationship between HE and health status indicators is controversial[20]. The reason to include a health status indicator is to identify whether there exist a correlation between expenditure and health level. **Life expectancy at birth** is an appropriate measure of indicator of health status for Canada. Life expectancy at birth

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[20] The studies showed that there is no correlation between HE and health status in the OECD countries (Kyriopoulos and Souliotis, 2002). However, a study in the United Kingdom provided evidence of correlation between the total health spending as a percentage of GDP and infant mortality rate (Maxwell, 1981).

simply gives the average number of years a new-born baby is expected to live. It is used as a proxy measure for health status, given the mortality risks[21]. However, the shortcoming of this variable is that it measures quantity rather than the quality of life[22]. A recent cross-sectional study of the determinants of health expenditures by Chern et al. (2002) using structural equation modeling revealed that health status and prior utilization of health services have considerable predictive power for future health spending.

A distinction should be made between the effects of variables that represent demographic structure and health status. An increasing share of senior population implies increasing health expenditures due to higher costs of treatment of the elderly. However, increasing life expectancy or health status is associated with long-term care. Theoretically or intuitively, the sign of this effect is ambiguous. If marginal increases in health status increase health expenditures, this would imply that more expenditure on health care is needed to make people live longer. However, if marginal increases in health status decrease health expenditures, the cost of maintaining previous levels of health decreases as the health condition of the society improves. This situation leads to less need for health care and thus less expenditure on health care.

Table 2.4 displays the life expectancy at birth for the period 1975-1999. Life expectancy at birth in Canada has increased by 4 years on average within the last 20 years, reaching almost 80 in 1999. The average provincial growth rate of life expectancy at birth ranges between 0.2 and 0.3 percent, showing very similar trends across provinces.

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[21] See World Health Organization website for a list of core indicators of health status.
http://www.euro.who.int/eprise/main/WHO/Progs/CHE/Monitoring/20020802_5?PrintView=1&
[22] See Statistics Canada (2003) *Health Indicators*, 82-221 XIE., no: 2

Figure 2.4: Life Expectancy at Birth (years)
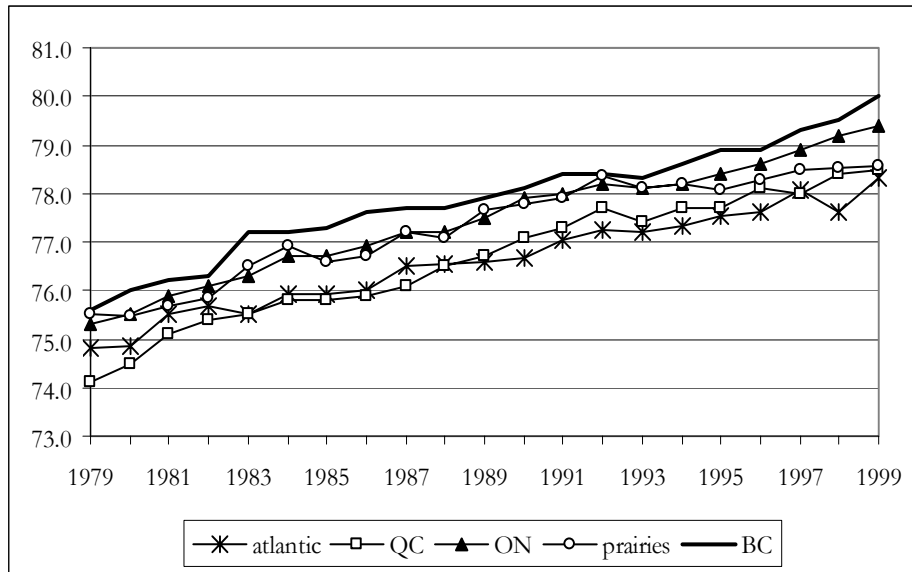
2.3.2 Preliminaries

This section presents the dynamics of provincial health expenditures. *One-way fixed effect error component models* are considered due to our focus on the provincial differences in health expenditures rather than differences across time. It is first assumed that these differences can be captured by the differences in the endowments.

The dynamic models considered are of such form:

$$h_{it} = \alpha + \rho h_{i,t-1} + X'_{it}\beta + \varepsilon_{it} \tag{2.1}$$

with $\quad \varepsilon_{it} = \mu_i + \upsilon_{it}$ (2.2)

i denotes the province, $i = 1,...,N$ and t denotes time, $t = 1,...,T$, β is a K x 1 vector where K is the number of explanatory variables, X_{it} is the it^{th} observation on K regressors, μ_i is the province specific effect and υ_{it} is the stochastic disturbance term. The following assumptions have been made:

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Assumptions

    i. h_{i0} is fixed.

    ii. $\mu_i \sim$ IID $(0, \sigma_\mu^2)$

    iii. $\upsilon_{it} \sim (0, \sigma_i^2)$

    iv. $E\left(X_{it}'\upsilon_{it}\right) \neq 0$

    v. $\upsilon_{it} = \varphi\upsilon_{i,t-1} + u_{it}$ $|\varphi| < 1$ and $u_{it} \sim$ IID $(0, \sigma_u^2)$

    vi. $E\left(h_{it}^2\right) < \infty; E\left(X_{it}^2\right) < \infty$

(i) is the initial condition that starts up the process in (2.1). This assumption is standard in the dynamic panel literature. (*ii*), (*iii*) and (*iv*) assume homoscedastic individual effects, cross-section heteroscedastic errors and failure of orthogonality or strict exogeneity for a subgroup of regressors respectively, (*v*) allows errors to follow an AR (1) process due to possible *through-time-allocated* effects of shocks to the error term[23] and (*vi*) assumes covariance stationarity. Further, the initial conditions are assumed to be fixed.

In vector form, (2.1) and (2.2) can be written as:

$$h = \alpha\iota_{NT} + h_{-1}\rho + X\beta + \varepsilon \tag{2.3}$$

$$\text{and } \varepsilon = Z_\mu\mu + \upsilon \tag{2.4}$$

$\varepsilon' = (\varepsilon_{11}, \ldots, \varepsilon_{1T}, \varepsilon_{21}, \ldots, \varepsilon_{2T}, \ldots, \varepsilon_{N1}, \ldots, \varepsilon_{NT})$ are the stacked errors, h is NT x 1, X is NT x K, ι_{NT} is NT x 1 vector of ones, Z_μ is NT x N matrix of individual dummies to be used in fixed effects estimation and $\mu' = (\mu_1, \ldots, \mu_N)$.

Substituting (2.4) into (2.3):

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[23] Ignoring homoscedastic and/or serially correlated errors yields consistent but inefficient estimates and biased standard errors.

$$h = \alpha \iota_{NT} + h_{-1}\rho + X\beta + Z_{\mu}\mu + \upsilon = Z\delta + Z_{\mu}\mu + \upsilon \qquad (2.5)$$

where $Z = \left[\iota_{NT}, X, h_{-1}\right]$ and $\delta^{'} = \left[\alpha^{'}, \rho^{'}, \beta^{'}\right]$

Equation (1) is treated under cross-section heteroscedasticity and contemporaneous correlation between residuals for province i and j such that:

$$E\left(\upsilon_{it}\upsilon_{jt}\middle|Z_t\right) = \sigma_{ij}$$

$$E\left(\upsilon_{is}\upsilon_{jt}\middle|Z_t\right) = 0 \qquad \text{for all } i, j, t, s \text{ with } t \neq s$$

The residual covariance is to be estimated using GLS weighting analogous to the covariance across cross-sections in a *seemingly unrelated regression* framework. In this case the covariance matrix can be written as:

$$\Omega = E\left[\upsilon\upsilon^{'}\right] = \begin{bmatrix} \sigma_{11}I_T & \sigma_{12}I_T & \cdots & \sigma_{1N}I_T \\ \sigma_{21}I_T & \sigma_{22}I_T & \cdots & \sigma_{2N}I_T \\ \vdots & \vdots & \ddots & \vdots \\ \sigma_{N1}I_T & \sigma_{N2}I_T & \cdots & \sigma_{NN}I_T \end{bmatrix} = \Sigma \otimes I_T$$

2.3.3 Generalized Instrumental Variable (TSLS) Estimation

Arellano (2003) demonstrated that the first problem encountered in dynamic panels is the correlation between the lagged dependent variable and the error term. Since h_{it} is correlated with the disturbance, it follows that $h_{i,t-1}$ will also be correlated with the disturbances through the error component even if the disturbances are not serially correlated. Therefore the lagged dependent variable is an endogenous variable. This will render biased and inconsistent OLS estimates. In fact the bias will be downward if $\rho > 0$. To overcome this problem, the estimation is done via Instrumental Variables (IV) where $h_{i,t-2}$ is uncorrelated with the error term and appropriate as an instrument for $h_{i,t-1}$. A second problem is the endogeneity of the explanatory variables

resulting in failure to satisfy the orthogonality condition which renders biased and inconsistent estimates. There are two variables in the dataset suspected to be endogenous in relation to health spending[24]. First, the relative price of health care can be thought as Granger caused by HE and therefore being predetermined rather than strictly exogenous. Second, life expectancy at birth as a proxy for health status is expected to be highly endogenous because changing health status may occur as a result of spending more on health care and vice-versa, suggesting that the causality may run from health spending to health status as well[25].

Returning to (2.5), one can obtain the generalized instrumental variable (GIV) estimator of δ as:

$$\hat{\delta}_{GIV} = \left(Z'\hat{\Omega}^{-1/2} P_W \hat{\Omega}^{-1/2} Z \right)^{-1} \left(Z'\hat{\Omega}^{-1/2} P_W \hat{\Omega}^{-1/2} h \right) \tag{2.6}$$

$P_W = (W(W'W)^{-1}W')$ is the projection matrix.

W_i^+, the matrix of instruments for province i is defined as (see Baltagi, 2001):

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[24] The endogeneity of income is not considered here and it is argued that causality runs from income to health spending and not the other way (see AC).

[25] The test for endogeneity initially proposed by Hausman (1978) has been generalized by Ahn and Low (1996, henceforth AL) to test the joint null hypothesis of exogeneity of the explanatory variables. To perform the AL test, the lagged dependent variable is dropped from the model to restrict our focus only on the joint exogeneity of the explanatory factors. After performing GLS on static (2.1) one can obtain the residuals, defined by $\varepsilon = (\sigma_\nu \Omega^{-1/2})h - (\sigma_\nu \Omega^{-1/2})X\hat{\beta}_{GLS}$. Then define $Q = I_{NT} - P = I_{NT} - Z_\mu (Z'_\mu Z_\mu)^{-1} Z'_\mu$ and $Z_\mu = I_N \otimes \iota_T$ where I_N is an N x N identity matrix, ι_T is a T x 1 vector of ones, $\otimes$ denotes Kronecker product, I_{NT} is NT x NT identity matrix and Z_μ is NT x N matrix. The residuals obtained from the GLS regression are regressed on $\widetilde{X}$ and X where $\widetilde{X} = QX$ and Q wipes out the individual effects. The LM test statistic is defined as $NTR^2 \sim \chi^2(k)$ under the null hypothesis where k is the number of explanatory variables.

$$W_i = \begin{bmatrix} \left[h_{i1}, x'_{2i1}, x'_{2i2}\right] & 0 & \cdots & 0 \\ 0 & \left[h_{i1}, h_{i2}, x'_{2i1}, x'_{2i2}, x'_{2i3}\right] & & \vdots \\ \vdots & 0 & \ddots & 0 \\ 0 & \cdots & 0 & \left[h_{i1}, \ldots, h_{i,T-2}, x'_{2i1}, \ldots, x'_{i,T-1}\right] \end{bmatrix}$$

$$W_i^{+} = \begin{bmatrix} W_i & & & & 0 \\ & \left[x'_{1i1}, x'_{1i2}\right] & & & \\ & & x'_{1i3} & & \\ & & & \ddots & \\ 0 & & & & x'_{1iT} \end{bmatrix}$$

where x_{1it} is strictly exogenous but x_{2it} is not.

$W = \left[W_1^{+}, \ldots, W_N^{+}\right]$ is the complete matrix of instruments satisfying $E\left[W'Z\right] \neq 0$; $E\left[W'\upsilon\right] = 0$.

2.3.4 Generalized Method of Moments Estimation

Anderson and Hsiao (1981) suggested that (2.5) can also be written in difference form to wipe out the individual effects. This method is further considered as a remedy against the above-mentioned drawback of OLS.

$$\Delta h = \Delta h_{-1}\rho + \Delta X\beta + \Delta\upsilon \tag{2.7}$$

The first differenced form introduces bias and serial correlation because the lagged dependent variable is correlated with the first order moving average error term. Therefore, the IV estimation is required to consistently estimate the parameters in (2.7), if not efficiently, where the appropriate instrument for $\Delta h_{i,t-1}$ is simply $h_{i,t-2}$[26].

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[26] Arellano (1989) suggests that using instruments in lagged level form ($\ln g_{i,t-2}$, $\ln g_{i,t-3}$,...) is preferable to using lagged differences ($\Delta \ln g_{i,t-2}$, $\Delta \ln g_{i,t-3}$,…) due to smaller coefficient standard errors.

However, (2.5) can be estimated efficiently via Generalized Method of Moments (GMM) using orthogonal deviations transformation.

The moment conditions are:

$$E\left(W_{it}^{'}\upsilon_{is}\right)=0 \qquad\qquad t=1,...,T \qquad t \leq s \qquad\qquad (2.8)$$

There are $T(T+1)k/2$ such moment conditions where k is the number of instruments. If the instruments are weakly exogenous with respect to ε_{it} but uncorrelated with the individual effects, then the moment conditions reduce to:

$$E\left(W_{it}^{'}\Delta\upsilon_{is}\right)=E\left(W_{it}^{'}\Delta\varepsilon_{is}\right)=0 \qquad\qquad t=1,...,T\text{-}1 \qquad t \leq s \qquad (2.9)$$

where $\Delta\upsilon_{is}=\upsilon_{is}-\upsilon_{i,s\text{-}1}$

If $\forall i$, ε_{it} is uncorrelated with h_{i0} and μ_{i} and if $\forall i$ the ε_{it} are mutually uncorrelated then the following moment conditions hold (see Mátyás, 1999):

$$E\left(h_{is}\Delta\upsilon_{it}\right)=0 \qquad\qquad t=2,...,T \qquad s=0,...,t-2 \qquad\qquad (2.10)$$

$$E\left(\upsilon_{iT}\Delta\upsilon_{it}\right)=0 \qquad\qquad t=2,...,T\text{-}1 \qquad\qquad\qquad\qquad (2.11)$$

where $\Delta\upsilon_{it}=\upsilon_{it}-\upsilon_{i,t\text{-}1}$

(2.10) and (2.11) imply $T(T-1)/2$ and $T-2$ such conditions respectively. There are additional moment conditions implied by the stationarity assumption which are not stated here. (see Arellano, 2003, p: 111)

The GMM estimator of (2.5) is[27]:

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[27] See Arellano (2003).

$$\begin{bmatrix} \hat{\rho} \\ \hat{\beta} \end{bmatrix} = \left([h_{-1}, X]' K' W A_N W' K [h_{-1}, X] \right)^{-1} \left([h_{-1}, X]' K' W A_N W' K h \right) \qquad (2.12)$$

where the weight matrix, A_N has to be chosen.

The one-step weight matrix is:

$$A_N = \left(W' K K' W \right)^{-1}$$

$K = (D \tilde{\Omega} D')^{-1/2} D$, $\tilde{\Omega}$ is the estimated covariance matrix of the errors in levels, D is the first difference operator.

$$D = \begin{bmatrix} -1 & 1 & 0 & \cdots & 0 & 0 \\ 0 & -1 & 1 & & 0 & 0 \\ \vdots & & & \ddots & & \vdots \\ 0 & 0 & 0 & \cdots & -1 & 1 \end{bmatrix}_{(T-1) \times T}$$

The two-step GMM uses the robust choice:

$$A_N = \left(W' K \tilde{\upsilon} \tilde{\upsilon}' K' W \right)^{-1} \qquad \text{where } \tilde{\upsilon} \text{ is the one-step residuals}$$

The nature of orthogonal deviation is to take first difference and apply GLS transformation to remove the artificial autocorrelation caused by first differencing[28]. Using orthogonal deviations in GMM estimation wipes out the individual effects as in first-differencing but does not introduce serial correlation in the transformed residuals. Analogous to 2-step, N-step GMM which will be used for the estimation continuously updates the weight matrix until convergence is achieved. The criterion function is:

$$Q_N(\delta) = f_N(\delta)' V(\delta)^{-1} f_N(\delta)$$

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[28] See Arellano (2003, p: 17) for technical discussion on this transformation.

where $f_N(\delta)$ are the moment conditions and $V(\delta)$ is the optimal weight matrix

The continuously updated GMM (CU-GMM) solves:

$$\hat{\delta} = \arg\min{}_{\delta} \, Q_N(\delta)$$

The advantage of 2-step or CU-GMM is that no prior assumption has to be made about the distribution of the errors and the individual heterogeneity (see Baltagi, 2001).

2.3.5 Models and Specification

Consider the following dynamic model for the government:

$$\ln g_{it} = (\alpha + \mu_i) + \beta_1 \ln y_{it} + \rho \ln g_{i,t-1} + \beta_2 \ln r_{it}$$
$$+ \beta_3 p65_{it} + \beta_4 x_{it} + \lambda t + \upsilon_{it} \qquad (2.13)$$

ln denotes the natural logarithm and the variable t denotes the linear time trend.

The inclusion of time trend has serious implications. First, health expenditures in Canada tend to increase over time therefore it may be appropriate to include a linear trend to separate its effect on the estimated long-run coefficients. Second, without a trend variable in (2.13), the t-statistics can be misleading due to the common trends. The linear trend can also be seen as a measure that captures the technological progress. Technology has an important role in the rising cost of health care (see Blomqvist and Carter, 1997).

From (2.13), the respective long-run income and price elasticity of government health expenditures are:

$$E_{g,y} = \frac{\beta_1}{1-\rho} \quad ; \qquad E_{g,r} = \frac{\beta_2}{1-\rho}$$

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The long-run effects of the share of senior population and life expectancy at birth respectively are:

$$\beta_3 = \frac{d \ln g}{dp65} = \frac{d \ln g}{dg} \times \frac{dg}{dp65} = \frac{1}{g} \times \frac{dg}{dp65} \qquad (2.14)$$

$$\beta_4 = \frac{d \ln g}{dx} = \frac{d \ln g}{dg} \times \frac{dg}{dx} = \frac{1}{g} \times \frac{dg}{dx} \qquad (2.15)$$

β_3 represents the relative change in g (i.e. $\Delta g/g$) resulting from a one-percentage point change in the share of senior population. Similarly, β_4 represents the relative change in g resulting from a one-year change in life expectancy at birth. If the interest is the percentage change in g resulting from a unit change in $p65$ and x than the estimated parameters of these variables should be multiplied by 100. The elasticity of government health expenditures with respect to share of senior population and life expectancy at birth can be obtained from (2.14) and (2.15) respectively:

$$E_{g,p65} = \frac{d \ln g}{d \ln p65} = \frac{dg}{dp65} \times \frac{p65}{g} = \beta_3 \times p65 \qquad (2.16)$$

$$E_{g,x} = \frac{d \ln g}{d \ln x} = \frac{dg}{dx} \times \frac{x}{g} = \beta_4 \times x \qquad (2.17)$$

Since $\beta_3 = \frac{1}{g} \times \frac{dg}{dp65}$ and $\beta_4 = \frac{1}{g} \times \frac{dg}{dx}$

Therefore, the elasticity for the share of senior population and life expectancy at birth can be obtained from (2.13). The advantage of writing the model in log-lin form

with respect to share of senior population and life expectancy at birth is that the elasticities can be evaluated at the point where it is relevant for policy.

Equation (2.13) can be written in such form that the estimated parameters are direct long-run elasticities or effects. This transformation is due to Bewley (1979). Subtracting $\rho \ln g_{i,t}$ on both sides, (2.13) becomes:

$$(1 - \rho)\ln g_{it} = (\alpha + \mu_i) + \beta_1 \ln y_{it} - \rho(\ln g_{it} - \ln g_{i,t-1}) + \beta_2 \ln r_{it}$$
$$+ \beta_3 p65_{it} + \beta_4 x_{it} + \lambda t + \upsilon_{it}$$

dividing by $(1 - \rho)$ gives:

$$\ln g_{it} = \left(\frac{\alpha + \mu_i}{1 - \rho}\right) + \left(\frac{\beta_1}{1 - \rho}\right)\ln y_{it} - \left(\frac{\rho}{1 - \rho}\right)\Delta \ln g_{it} + \left(\frac{\beta_2}{1 - \rho}\right)\ln r_{it}$$
$$+ \left(\frac{\beta_3}{1 - \rho}\right)p65_{it} + \left(\frac{\beta_4}{1 - \rho}\right)x_{it} + \left(\frac{\lambda}{1 - \rho}\right)t + \left(\frac{1}{1 - \rho}\right)\upsilon_{it} \qquad (2.18)$$

The constant term in (2.18) can be seen as the steady-state mean for province i, if we let T go large. This transformation also requires IV estimation due to the correlation between transformed lagged dependent variable and the error term where $\ln g_{i,t-2}$ is the appropriate instrument for $\Delta \ln g_{it}$. The remaining regressors in (2.18) can serve as their own instruments as long as they are strictly exogenous. However, it is expected that this is not the case. The LM test statistic for the AL test of joint exogeneity of the explanatory variables for the static version of (2.13) turned out to be far greater than $\chi^2 (5)$, resulting in rejection of the null hypothesis of joint exogeneity. In this case, either lags of endogenous variables or external measures, given they are valid, can be used as instruments as long as they satisfy the orthogonality condition. However, finding valid instruments is an empirical problem that researchers should be extremely careful about. Two-period lagged value of relative prices and five period

lagged value of life expectancy are used as instruments to avoid correlation between the instruments and the error term of HE due to possible long-memory. An instrument is said to be weak if it fails to satisfy the orthogonality or if the relative bias of IV compared to OLS is above an arbitrary threshold tolerance level[29].

Reconsider (2.5):

$$h = Z\delta + \varepsilon \tag{2.19}$$

$$Z = W\Pi + v \tag{2.20}$$

In technical terms, given the reduced form for Z the IV estimator becomes more biased as the concentration parameter, $C = \Pi'Z'Z\Pi / \sigma_{vv}$, gets smaller. This paper is rather more concerned about the correlation between the instruments and the error term of the HE equation because any potential instrument is more likely to be correlated with the error term than its failure to be correlated with the endogenous regressors[30].

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[29] Bound et al. (1995) draw attention to the biases in IV estimator when the instruments are weak and argues that the IV bias decreases as the R^2 between the instruments and the endogenous regressor increases.

[30] There is a growing literature on detecting weak instruments in IV estimation of which the test proposed by Stock and Yogo (2001) is considered here. If Y, X and Z denote the endogenous, exogenous and excluded instrument matrices in stacked form respectively and K_1 and K_2 denote the number of exogenous and the number of excluded instruments respectively the Stock-Yogo test is based on the minimum eigenvalue of the concentration matrix:

$$g_{min} = \min eigenvalue(G_T)$$

$$G_T = \hat{\Sigma}_{vv}^{-1/2'} Y^{\perp'} P_{Z^\perp} Y^\perp \hat{\Sigma}_{vv}^{-1/2} / K_2$$

where $\hat{\Sigma}_{vv} = (Y'M_{\bar{Z}}Y)/(T-K_1-K_2)$, $Y^\perp = M_X Y = (I - X(X'X)^{-1}X')Y$,

$$M_{\bar{Z}} = I - XZ((XZ)'XZ)^{-1}(XZ)', \ Z^\perp = M_X Z = (I - X(X'X)^{-1}X')Z, P_{Z^\perp} = Z^\perp(Z^{\perp'}Z^\perp)^{-1}Z^{\perp'}$$

The critical values of this test statistic are tabulated by simulation and they depend on the relative bias tolerance and the number of endogenous regressors. A g_{min} above the critical value implies rejection of null of weak instruments. Results of the Stock-Yogo (SY) test fail to reject the null hypothesis of weak instruments suggesting that there may be a certain degree of correlation between the instruments and the error term of HE equation. However, the implementation of this test in panel context is not clear since it is derived under time-series properties and there are additional assumptions related to the individual heterogeneity whose effect on the test statistic is unknown. Therefore the SY test results should be treated with extreme caution.

CHAPTER 3

RESULTS & POLICY IMPLICATIONS

The assumption made about the differences in the intercepts has been incorporated in the models via fixed effects[31]. The results show that the dynamics of HE should not be ignored as they play a significant role in the adjustment process of explanatory variables. An interesting result found is that the life expectancy at birth has statistically significant effect on government HE. Before analyzing the precise effects of those variables, we should confine ourselves to the reparameterized models we made use of, based on Bewley (1979), to directly estimate the average long-run effects of the explanatory variables. This reparameterization helps to assess the significance of the long-run effects and their standard errors. Table 4 reports the results. All factors have statistically significant long-run effects on total HE with the exception of the share of senior population. The long-run price elasticity of total health expenditure is statistically significant at 5% and carried a negative sign suggesting that a 1% increase in relative prices decreases total health spending by 0.12% on average[32]. The long-run income elasticity of total health expenditure is 0.36 suggesting that a 1% increase in per capita GDP is associated with a slower increase of total health expenditure around 0.36%. Consistent with the theory, the effect of the share of publicly funded health expenditure on total health expenditures carried a positive sign. But the magnitude of this effect is small.

Concerning the government HE, all long-run effects are significant. The evidence suggests that the effect of the share of senior population is neither high as it

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[31] We have performed F-test to test the joint significance of the individual fixed effects under the null hypothesis, H_o: $\mu_1 = \mu_2 = \ldots = \mu_{10} = 0$. The F-test turned out to be 41.81, 22.85 and 131.70 for total, government and private HE respectively, resulting in favor of rejecting the null hypothesis. Therefore, the models can be characterized by allowing fixed effects.

[32] In his seminal work, Newhouse (1977) argued that price may not be an important factor in explaining health expenditures if health care are heavily subsidized.

is previously found by DD[33], nor insignificant as argued by AC. The long-run income elasticity of government health expenditure is found to be 0.43. The evidence in this paper also indicates that the long-run effect of relative price of health care is more pronounced for the government with a price elasticity of -0.74. A possible explanation of the significance of price effect is that provincial governments face the full price of health services even though the cost is not projected on patients through billings. Regardless of this fact, the provision of public health care is not free and there are national constraints and long-term issues in financing of public health spending (see Brown, 1991).

The health status effect on government spending is considerably large. For the sample period, the government health expenditures decrease by 19% as result of a one year increase in life expectancy[34]. If it can roughly be postulated that life expectancy at birth increases by one year on average in every 7 years in Canada, this result indicates a considerable shrinkage in government health expenditures, given other factors constant. This finding indicates that improvement in health status leads to less need and thus less use of health care. On the other hand, increasing life expectancy also implies a greater share of senior population in the long-run. As the regression results in Table 4 show, the negative effect of health status outweighs the positive effect of the aging population on government health expenditures indicating that despite the upward movement in government health expenditures due to increasing share of the elderly, improvements in health status bring net gains in terms of less

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[33] According to DD, the impact of the log of the share of the population over the age of 65 on log of government health expenditures is found to be 0.81 whereas AC found no evidence on its statistical significance. It should be noted that DD has chosen a log-log functional form with respect to *p65* whereas a log-linear functional form has been chosen in this study with respect to *p65*.

[34] It proved impossible to properly instrument the life expectancy at birth by its lags in the Total HE model. Therefore, it is thought appropriate in this situation to drop this variable from that model leaving its effect to remain unknown on total health spending.

spending[35]. This outweighing is one of the reasons for decreasing government health expenditures as a result of increasing life expectancy.

The estimation results for private health expenditures are reported in the last column of Table 4. The long-run income elasticity of private health expenditures is found to be 0.26, lower than that of government HE. This indicates that government health expenditures increase faster than the private health expenditures and thus a 1% increase in per capita GDP is associated with increasing centralization[36], *ceteris paribus*. This result contradicts the findings of Di Matteo (2000). For the relative prices, the long-run price elasticity is statistically insignificant. A possible reason for the insignificance of the price effect is that many consumers of private health care do not directly face full prices because of private insurance. Also in face of high private insurance coverage, income elasticities are close to zero (see Getzen, 2000). The empirical evidence presented here also confirms this argument.

The share of publicly funded HE is included in the analysis of the private sector to evaluate a potential trade-off between private and public health expenditures and its size. The empirical evidence contradicts the *a priori* expectation. Our findings indicate a statistically significant but negative trade-off between the share of public HE and private HE.

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[35] The outweighing will always hold not matter what the functional form of the model is with respect to share of senior population and life expectancy at birth as long as the coefficients of these variables have the same interpretation. To show this, reconsider (2.16) and (2.17):

$$\mathrm{E}_{g,p65} = \beta_3 \times p65$$
$$\mathrm{E}_{g,x} = \beta_4 \times x$$

From (2.18), the results in Table 4 for the government health expenditures model show that:

$$\left|\hat{\beta}_4\right| > \left|\hat{\beta}_3\right|$$

For the sample period of 1975-2002, since the values of life expectancy at birth and share of senior population does not overlap, $x_t > p65_s$ for $\forall t \in [1975 - 2002]$ and $\forall s \in [1975 - 2002]$. It follows that:

$$\left|\beta_4 \times x\right| = \left|\mathrm{E}_{g,x}\right| > \left|\mathrm{E}_{g,p65}\right| = \left|\beta_3 \times p65\right| \qquad \text{for } \forall t \in [1975 - 2002]$$

Even if the model is estimated such that the coefficients of the share of senior population and life expectancy are elasticities, the effect of health status would still outweigh the effect of the share of senior population.

[36] Centralization is defined as the ratio of public health expenditures to total health expenditures (see Di Matteo, 2000, p: 97).

The coefficient of trend appeared to be significant at 1% level for all three models. This suggests that if the remaining explanatory factors are fixed, total, government and private health expenditures would grow at a rate of 2% and 3.5% and 1.6% respectively. This result may have occurred as a consequence of fast-growing cost of medical technology.

The last homogenous model is estimated via GMM. It should be noted that the GMM estimation is designed and expected to perform well under large N small T environments which is not our case. The preliminary models with a full set of explanatory variables are first estimated and then the variables with insignificant parameters[37] are dropped from the model in order to obtain parsimonious models and to increase the sample size because not all of the explanatory variables have the same time span. The resulting models are displayed in Table 5. The GMM estimates were not subject to Bewley transformation therefore the parameters should be multiplied by the long-run multiplier, $1/(1 - \rho)$ in order to obtain the long-run effects.

The long-run income elasticity of total health spending turned out to be 0.35, very close to GIV estimate of 0.36. The effect of relative price of health care on total health spending in the GMM estimation is significant at 10% level with a price elasticity of -0.25 compared to -0.12 obtained via GIV. For the government, the GMM estimates a higher income elasticity of 0.52 and price elasticity of -0.52 compared to their respective GIV counterparts which are 0.43 and -0.74. However, the effects of the senior share and life expectancy are not significantly different from zero in GMM estimation. For the private sector, the GMM long-run effects of income and public provision are slightly higher than their GIV counterparts. The GMM long-run income elasticity is 0.31 and the effect of the share of publicly funded health expenditure is -0.038. Autonomous growth in health spending is lower in GMM compared to GIV estimates with 1.7%, 1.3% and 2% for total, government and private health expenditures respectively. A summary of comparative results of GMM and GIV

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[37] These variables are the share of senior population and life expectancy at birth.

estimates is given in Table 6. It should be noted that both GIV and GMM estimates are unbiased and consistent. However, only GMM is efficient.

The evidence in this paper confirms that health is far from being a luxury for Canada and the delivery of health is dominated by the needs rather than the ability to pay. This is what Culyer (1988) was referring to as "*the Bioengineering view*". The "*according to needs*" argument is also consistent with physicians having a high-degree of control over the decisions about the medical services that patients need.

CHAPTER 4

CONCLUSION

This thesis focused on the magnitude of income elasticity and the impact of non-income determinants of health expenditures in Canadian provinces using panel data on per capita GDP, relative price of health care, the share of publicly funded health expenditures, share of senior population and the life expectancy at birth over the period 1975-2002. Under the assumption of homogenous parameters, the income elasticity of health expenditure is below unity. This result is consistent with previous studies in the sense that the regional estimates are usually below one.

The main differences captured in this study are summarized as follows:

- The relation between health expenditure and its determinants is of autoregressive structure.
- Government health expenditures are constrained by the relative prices.
- Statistically significant effects of the share of senior population and share of publicly funded health expenditures are small.
- There is correlation between health spending and health status.

The inclusion of time trend for statistical reasons and to capture possible changes in medical technology is of importance since the results provide evidence of considerable autonomous growth of health spending, *ceteris paribus*. This paper also supports the findings of Blomqvist and Carter (1997) that when the time trend is excluded from the regressions, the long-run income elasticity is higher.

The difficulty encountered in this paper was indecision whether or not the panel can be described as group stationary. The IPS and Hadri's panel unit root tests gave contradictory result regarding the unit root problem. Most of the panel unit root tests

are based on and therefore valid only under joint or sequential limit theory and evidence presented confirms that these tests are known to render conflicting results due to their high/low power in certain cases. Based on this problem, it is argued that the effects of shocks to Canadian public sector can be best characterized as temporary rather than permanent and there is no firm reason and evidence to not to follow a traditional analysis. However, a thorough and careful assessment of panel unit root problem is needed.

A second area for future research lies in more advanced econometric techniques to reconsider the soundness of macroeconomic health policies regarding health expenditure. Cointegration approach is not followed here due to the unknown order of integration of the variables and short time span but one can expect to have a relationship at least between health spending, income and the relative price of health care in the long-run. After a careful assessment of unit root the cointegrating relationships can be examined to identify the long-run effects and the short-run dynamics.

Some of the studies of health care expenditure based on the OECD health data argued that there are substantial differences in the structure of health sectors and demographics in the OECD countries. It is also argued that imposing slope homogeneity is unrealistic and may lead to misleading coefficients (Roberts, 1999). Slope heterogeneity is not considered here due to well known misleading and biased estimates under dynamic pooled estimation with heterogeneous parameters (Pesaran and Smith, 1995). Also the preliminary estimation under heterogeneity gave conflicting results.

Extreme caution should be exercised when interpreting the results. The small sample behavior of GMM and the validity of the instruments are questionable matters and they indicate that some bias may have not been removed. A difficult task was to find a set of "clean" instruments for the variables suspected of being non-exogenous. Using lags of endogenous variables as instruments is a standard "remedy" in situations where it is almost impossible to find other "clean" instruments. Using sufficiently

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lagged values to make them clean results in loss of sample size equal N times the number of maximum lagged periods used in the estimation. These trade-offs should be taken into account when working with single equations and adjustments should be made according to which type of bias one would prefer to tolerate. A fruitful methodology for future research this study suggests is to implement a SUR methodology in system of equations considering the difficulties of finding valid instruments under single equation models and the distinct characteristics of the public and private equations one might have in the system. The M-equation framework is known to accommodate the endogeneity of the regressors which becomes the crucial part of the analysis if one focuses on the qualitative determinants[38] of health spending.

What are not needed are further studies of the effects of quantitative measures on health expenditures. The standard measures thought to be the determinants of health expenditures are so far known to researchers in this area. Further, the empirical evidence showed that these quantitative measures increase at a slower rate than health expenditures and thus they help very little in explaining the rise of health expenditures. As Culyer (1988) pointed out, beyond the questionability of econometric procedures and the statistical plausibility there is something systematically incorrect and inadequate with the quantitative analysis of health spending. This issue brings us to what is known as the *"ad hoc"[39]* feature of the analysis of health expenditures. *Ad hoc* models raise difficulties in setting sound policies especially in the developed countries where there are numerous empirical studies but they are limited to the question of whether health care is a necessity or a luxury good.

What is not known is the precise effect of measures that are indicators of the quality of life and health. Proxy measures for the effectiveness of the health system, the quality of health services and health status can serve for such purposes, thereby allowing one to examine the consequences of an increase in the quality of health care

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[38] Some of those proxy measures are the quality adjusted life years (QALY's), waiting time for an operation, length of stay in hospital, number of post-operation complications.

[39] The right-hand side variables are chosen intuitively or "arbitrarily" and are thought to affect health spending.

on health expenditures. Therefore, the next generation of international or regional comparisons of health spending should be based on the effects of qualitative measures that are truly responsible for the persistent increase or disparities in health expenditures.

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Table 1: Estimates of Income Elasticity of Health Expenditure

Year	Author	Type of Analysis	Data Source	Methodology	Variables	Sample	Income elasticity
1977	Newhouse	Cross-section	United Nations	Linear, log-linear, reciprocal forms	GDP	13	(>1)
1992	Gerdtham et al.	Cross-Section	OECD	Logarithmic functional form	Income, relative price of health, age, urbanization, public provision & financing, share of public HE	19	(>1)
1992	Gbesemete & Gerdtham	Cross-Section	African countries	Logarithmic functional form	GNP, foreign aid, crude birth rate, share of population under 15, urbanization rate	30	Around 1
1997	Blomqvist & Carter	Panel(D)	OECD	Nonstationarity Cointegration	GDP, share of senior population	744	(>1)
1998	Casasnovas & Saez	Panel (S)	OECD	Random effects, heterogeneity	GDP, share of senior population, share of public HE	880	(<1)
1998	Dimatteo & Dimatteo	Panel (S)	Canada	Fixed effects model	GDP, federal transfers, share of senior population	270	(<1)
1999	Roberts	Data Field (D, S)	OECD	Nonstationarity Heterogeneity Sensitivity Analysis	Publicly funded HE, GDP, price of health, Share of senior population	660	Around 1
1999	Hitiris	Panel (S)	G7 countries	Fixed effects model, linear functional form	GDP, Share of senior population, share of public expenditure in GDP	217	Around 1
2001	Okunade & Karakus	Panel (S)	OECD	Nonstationarity Cointegration, ECM	GDP, price of health	722	(>1)
2001	Ariste & Carr	Panel (S)	Canada	Nonstationarity Cointegration Fixed effects	GDP, Share of senior population, ratio of deficit to GDP	330	(<1)
2002	Clemente et al.	Panel (S)	EU & OECD	Heterogeneity Nonstationarity Cointegration Structural Break	GDP	185/333/296	(>1)
2002	Bac & Le Pen	Panel (S)	OECD	Nonstationarity Cointegration Fixed effects	GDP, price of health care	432	(>1)

Note: "D" and "S" denote dynamic and static models respectively.

Table 2: Province by Province ADF τ-statistics and IPS Panel t-bar statistic

Province	Total Health Expenditure		Government Health Expenditure		Private Health Expenditure		Share Public Health Expenditure	
	Lag order	τ-statistic	Lag order	τ-statistic	Lag order	τ-statistic	Lag order	τ-statistic
Newfoundland	2	-1.841	3	-1.859	4	-1.939	3	-2.828
Prince Edward Island	2	-3.410***	4	-1.740	2	-3.586***	2	-3.395***
Nova Scotia	2	-2.012	3	-2.126	3	-2.176	3	-1.974
New Brunswick	3	-3.419***	3	-2.821	2	-4.062**	3	-3.739**
Quebec	4	-2.551	1	-2.547	2	-2.207	2	-2.647
Ontario	2	-2.391	3	-2.361	3	-2.553	2	-1.538
Manitoba	3	-2.870	3	-2.947	3	-2.467	2	-2.255
Saskatchewan	2	-1.774	3	-2.882	2	-3.726**	3	-2.851
Alberta	2	-2.442	2	-2.417	3	-2.778	2	-2.178
British Columbia	4	-3.832**	3	-3.201	2	-2.067	2	-3.632**
Panel t – bar statistic		-2.654**		-2.539***		-2.756**		-2.713**

Note: ADF regressions include linear trend. *, ** and *** represent 1%, 5% and 10% significance levels respectively. The 1%, 5% and 10% critical values of the IPS t-bar test statistic are -2.79, -2.60 and -2.51 respectively.

Table 2: Province by Province ADF τ-statistics and IPS Panel t-bar statistic (cont'd)

Province	GDP		Relative Price of Health ☆		Life Expectancy at Birth		Share of Senior Population ☆	
	Lag order	τ-statistic	Lag order	τ-statistic	Lag order	τ-statistic	Lag order	τ-statistic
Newfoundland	3	-2.364	3	-3.956*	2	-4.031**	1	-0.052
Prince Edward Island	0	-3.516***	3	-4.402*	1	-3.486***	2	-1.711
Nova Scotia	3	-1.723	1	-4.767*	1	-3.041	2	-3.904*
New Brunswick	0	-3.071	1	-2.101	1	-2.380	2	-2.921
Quebec	2	-2.547	1	-1.264	1	-2.864	2	-1.667
Ontario	2	-2.900	2	-1.796	1	-2.776	3	-2.749
Manitoba	3	-2.347	1	-1.522	2	-1.805	3	-2.904
Saskatchewan	2	-1.186	4	-1.852	1	-1.572	3	-1.521
Alberta	1	-2.078	2	-2.242	3	-3.220	2	-0.857
British Columbia	3	-4.082**	2	-1.892	3	-3.480***	3	-1.619
Panel t – bar statistic		-2.582***		-2.579*		-2.865*		-1.9905**

Note: ☆ represents that the ADF regressions do not include linear trend. *, ** and *** represent 1%, 5% and 10% significance levels respectively. The 1%, 5% and 10% critical values of the IPS t-bar test statistic are -2.21, -1.99 and -1.89 respectively.

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Table 3: Province by Province KPSS η-statistics and Hadri's Panel Test statistic

Province	Total Health Expenditure $l_4 = 3$		Gov. Health Expenditure $l_4 = 3$		Private Health Expenditure $l_4 = 3$		Share of Public Health Expenditure $l_4 = 3$	
	η_τ	η_μ	η_τ	η_μ	η_τ	η_μ	η_τ	η_μ
Newfoundland	0.112	0.780**	0.089	0.779**	0.112	0.219	0.087	0.527**
Prince Edward Island	0.066	0.790**	0.074	0.771**	0.066	0.675**	0.067	0.078
Nova Scotia	0.140	0.767**	0.125	0.741**	0.131	0.786**	0.094	0.740**
New Brunswick	0.173**	0.771**	0.161**	0.765**	0.158**	0.760**	0.100	0.648**
Quebec	0.098	0.791**	0.113	0.745**	0.095	0.776**	0.098	0.746**
Ontario	0.155**	0.774**	0.151**	0.697**	0.126	0.806**	0.143	0.690**
Manitoba	0.115	0.776**	0.108	0.734**	0.091	0.771**	0.097	0.465**
Saskatchewan	0.132	0.762**	0.133	0.686**	0.091	0.747**	0.124	0.215
Alberta	0.128	0.672**	0.130	0.463	0.057	0.794**	0.143	0.624**
British Columbia	0.100	0.795**	0.091	0.778**	0.054	0.794**	0.085	0.364
Hadri Panel Statistic	4.973**	12.78**	4.516**	11.88**	2.973**	12.40**	2.297**	8.406**

Note: η_τ and η_μ are the trend and the level stationarity cases respectively. The 5% critical value of the Hadri Panel statistic is 1.645. ** denotes 5% significance level.

Table 3: Province by Province KPSS η-statistics and Hadri's Panel Test statistic (cont'd)

Province	GDP $l_4 = 3$		Relative Price of Health $l_4 = 3$		Life Expectancy at Birth $l_4 = 3$		Share of senior population $l_4 = 3$	
	η_τ	η_μ	η_τ	η_μ	η_τ	η_μ	η_τ	η_μ
Newfoundland	0.089	0.677**	0.181**	0.534**	0.088	0.624**	0.098	0.805**
Prince Edward Island	0.110	0.660**	0.171**	0.615**	0.101	0.604**	0.185**	0.737**
Nova Scotia	0.167**	0.625**	0.182**	0.576**	0.133	0.643**	0.207**	0.791**
New Brunswick	0.113	0.652**	0.123	0.633**	0.166**	0.626**	0.202**	0.796**
Quebec	0.076	0.628**	0.101	0.717**	0.127	0.643**	0.153**	0.807**
Ontario	0.066	0.588**	0.139	0.709**	0.117	0.644**	0.193**	0.798**
Manitoba	0.106	0.597**	0.126	0.597**	0.153**	0.605**	0.205**	0.751**
Saskatchewan	0.158**	0.450	0.170**	0.585**	0.140	0.585**	0.147**	0.773**
Alberta	0.122	0.223	0.177**	0.566**	0.145	0.632**	0.122	0.763**
British Columbia	0.046	0.637**	0.145	0.459	0.110	0.643**	0.185**	0.749**
Hadri Panel Statistic	4.051**	11.87**	7.932**	9.01**	4.68**	9.72**	8.43**	13.14**

Note: η_τ and η_μ are the trend and the level stationarity cases respectively. The 5% critical value of the Hadri Panel statistic is 1.645. ** denotes 5% significance level.

Table 4: Direct Long-run Estimates, 1979 – 1999
Method: Instrumental Variables, one-way fixed effects error component model

Variable	Total Health Expenditures		Government Health Expenditures		Private Health Expenditures	
	Coefficient (s.e)	p-value	Coefficient (s.e)	p-value	Coefficient (s.e)	p-value
GDP	0.36 (0.061)	0.0000	0.43 (0.153)	0.0049	0.26 (0.059)	0.0000
Price of Health Care	-0.12 (0.056)	0.0255	-0.74 (0.195)	0.0002	0.008 (0.091)	0.9226
Share of Public H.E	0.003 (0.001)	0.0133	-		-0.034 (0.001)	0.0000
Share of Senior Population	-0.004 (0.008)	0.6086	0.080 (0.016)	0.0000	0.027 (0.008)	0.0019
Life Expectancy at Birth	-		-0.191 (0.068)	0.0060	-0.015 (0.030)	0.6142
Trend	0.02 (0.001)	0.0000	0.035 (0.012)	0.0054	0.016 (0.004)	0.0003
Change in Dependent Var.	-1.87 (0.312)	0.0000	-1.74 (0.431)	0.0001	-0.35 (0.076)	0.0000
Constants						
Newfoundland	3.5527	0.0000	16.3453	0.0002	6.8802	0.0265
Prince Edward Island	3.6340	0.0000	16.1198	0.0002	6.8817	0.0261
Nova Scotia	3.5555	0.0000	16.0712	0.0002	6.8432	0.0271
New Brunswick	3.5773	0.0000	16.1782	0.0002	6.8932	0.0265
Quebec	3.5099	0.0000	16.2270	0.0002	6.8814	0.0279
Ontario	3.5389	0.0000	16.2770	0.0003	6.9214	0.0284
Manitoba	3.5971	0.0000	16.1263	0.0003	6.8748	0.0273
Saskatchewan	3.4738	0.0000	16.0599	0.0003	6.7872	0.0294
Alberta	3.4033	0.0000	16.3556	0.0003	6.9083	0.0299
British Columbia	3.5726	0.0000	16.3759	0.0003	6.9100	0.0289
Sample Size = 230/160/160	$R^2 = 0.99$		$R^2 = 0.99$		$R^2 = 0.99$	
	J-Statistic ≈ 0		J-Statistic ≈ 0		J-Statistic ≈ 0	
	Jarque Bera = 2.999 (0.2231)		Jarque Bera = 0.515 (0.7728)		Jarque Bera = 0.462 (0.7936)	

Note: Two-period lagged value is used as instrument for the lagged dependent variable and the relative price of health care. Five-period lagged value is used as an instrument for the life expectancy. Standard errors in parentheses are robust to heteroscedasticity of any form.

Table 5: Generalized Method of Moments Estimates, 1978 – 2002

Variable	Total Health Expenditures		Government Health Expenditures		Private Health Expenditures	
	Coefficient (s.e)	p-value	Coefficient (s.e)	p-value	Coefficient (s.e)	p-value
GDP	0.10 (0.040)	0.0107	0.12 (0.044)	0.0047	0.21 (0.087)	0.0140
Price of Health Care	-0.07 (0.038)	0.0659	-0.12 (0.042)	0.0038	-	
Share of Public H.E	-		-		-0.026 (0.005)	0.0000
Trend	0.005 (0.001)	0.0000	0.003 (0.001)	0.0002	0.014 (0.002)	0.0000
Lagged Dependent Variable	0.72 (0.048)	0.0000	0.77 (0.037)	0.0000	0.33 (0.100)	0.0011
Sample Size = 220/220/250	Adj. R^2 = 0.85 J-Statistic ≈ 0 Jarque Bera = 6.358 (0.0416)		Adj. R^2 = 0.77 J-Statistic ≈ 0 Jarque Bera = 0.318 (0.8528)		Adj. R^2 = 0.91 J-Statistic ≈ 0 Jarque Bera = 0.098 (0.9518)	

Note: Two-period lags are used as instrument for the lagged dependent variable and relative price of health care. Exogenous explanatory variables served as their own instruments. The transformation is via orthogonal deviations, the GMM weights are Arellano & Bond n-step period weights and the standard errors in parentheses are robust to period heteroscedasticity and serial correlation.

Table 6: Comparative GIV and GMM results

Variable	Total Health Spending		Gov. Health Spending		Private Health Spending	
	GIV	CU-GMM	GIV	CU-GMM	GIV	CU-GMM
Income	0.36	0.35	0.43	0.52	0.26	0.31
Relative Price	-0.12	-0.25	-0.74	-0.52	insignificant	insignificant
Public Share	0.003	insignificant	-	-	-0.034	-0.038
Senior Share	insignificant	insignificant	0.080	insignificant	0.027	insignificant
Health Status	-	insignificant	-0.19	insignificant	insignificant	insignificant
Trend	0.020	0.017	0.035	0.013	0.016	0.020

APPENDIX

A.1 Test of Null of Unit Root

The Augmented Dickey-Fuller test can be shown by the following model:

$$\Delta x_{it} = \alpha_i + \delta_i t + (1 - \rho_i) x_{i,t-1} + \sum_{j=1}^{K_i} \beta_{i,j} \Delta x_{i,t-j} + \varepsilon_{it} \qquad \varepsilon_{it} \sim i.i.d\,(0, \sigma^2)$$

where variable t is time trend for the ith province, $t = 1,......,T$ and $j = 1,......,K$. K is the number of lags, determined such that the error term is autocorrelation free.

IPS proposed a panel unit root test based on the average of the ADF test statistics:

$$\bar{t}_{NT} = \frac{1}{N} \sum_{i=1}^{N} \tau_i \qquad i = 1,.....,10$$

where τ_i is the ADF test statistic for i^{th} province.

The t-bar statistic converges in probability to a standard normal variate as $T \to \infty$ and $N \to \infty$. The null hypothesis that all series contain unit roots is tested against the alternative that some series are stationary.

$H_o: \rho_i = 1$ for all i

$H_A: \rho_i < 1$ $i = 1, 2, ..., N_1$ where N_1 is a subset of N

Based on the theorem 3.1 in Im et al. (2003), the standardized t-bar statistic is given by:

$$Z_{\bar{t}} = \frac{\sqrt{N}\{\bar{t}_{NT} - E(\tau_T)\}}{\sqrt{\text{var}(\tau_T)}}$$

$E(\tau_T)$ and var(τ_T) are respectively the mean and the variance of the individual ADF test statistic τ_i. The construction of the Z-statistic assumes that the individual ADF test statistics are iid with finite moments. However, the Z-statistic requires that, $N \to \infty$ and therefore cannot be used in this case. Instead, they propose a fixed T, fixed N test which is the t-bar statistic. But in this case, the statistic has a non-normal distribution and the critical values are supplied by IPS. A particular lag order is determined for each of the series instead of choosing a common lag order to avoid misleading ADF statistics resulting from autocorrelation.

A.2 Test of Null of Stationarity

The KPSS unit root test constructs the null hypothesis of stationarity against the alternative of unit root. This ensures that the null will be rejected only when there is strong evidence against it. Due to Kwiatkowski et al. (1992), a time series can be decomposed into three components, a deterministic trend, a random walk and a stationary error:

$$x_{i,t} = \theta_i t + r_{i,t} + \varepsilon_{i,t} \tag{1}$$

where t captures the deterministic trend and $r_{i,t}$ is a random walk:

$$r_{i,t} = r_{i,t-1} + u_{i,t} \qquad\qquad u_{i,t} \sim i.i.d\,(0, \sigma_u^2) \tag{2}$$

To test the null hypothesis of trend stationarity, the estimate of the error variance, est. $\sigma_{\varepsilon,i}^2$ and the residuals, $\varepsilon_{i,t}$ can be obtained by regressing x on a constant and time trend as in eq. (1). The same argument can be applied to test the null of level stationarity. In that case, x is regressed on a constant only. The test statistic is a one-sided LM statistic under the null of level stationary (H_o: $\theta_i = 0$) with the errors being *iid* in eq. (1). The LM test statistic is defined as:

$$\eta_i = \frac{1}{T^2} \sum_{t=1}^{T} S_{i,t}^2 / \hat{\sigma}_{\varepsilon,i}^2(l)$$

where T is the sample size, $\hat{\sigma}_i^2(l)$ is the estimate of the error variance, l is the lag truncation parameter[40] and $S_{i,t}$ is the partial sums of the residuals, $S_{i,t} = \sum_{j=1}^{t} \hat{\varepsilon}_{i,j}$. Instead of adding lagged terms into the unit root regression under testing as in the ADF, the KPPS test makes a nonparametric correction of the estimate of the error variance such that:

$$\hat{\sigma}_{\varepsilon,i}^2(l) = \frac{1}{T} \sum_{t=1}^{T} \varepsilon_{i,t}^2 + \frac{2}{T} \sum_{s=1}^{l} \left(\frac{1-2}{1+l} \right) \sum_{t=s+1}^{T} \hat{\varepsilon}_{i,t} \hat{\varepsilon}_{i,t-s}$$

The extension of the KPSS test for panel data has been proposed by Hadri (2000). The panel LM test statistic is defined as the mean of the individual test statistics under the null of level stationary:

$$L\hat{M}_\mu = \frac{1}{N} \sum_{i=1}^{N} \eta_i$$

The null hypothesis of level or trend stationarity is tested against the alternative of unit root in panel. Under the assumptions that $E[u_{i,t}] = E[\varepsilon_{i,t}] = 0$, $u_{i,t}$ and $\varepsilon_{i,t}$ are *i.i.d* across i and over t, the test statistic has the following limiting distribution:

$$Z_\mu = \frac{\sqrt{N}(L\hat{M}_\mu - \xi_\mu)}{\zeta_\mu} \Rightarrow N(0,1)$$

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[40] Lag truncation is set to *integer* $[4(T/100)^{1/4}]$ to correct the estimate of the error variance.

where $\Rightarrow$ represents weak convergence in distribution, ξ_μ, ζ_μ are mean and variance of a standard Brownian bridge $\int_0^1 V^2(r)dr$. The computed numerical values of ξ_μ, ζ_μ are 1/6 and 1/45 for the level case and 1/15 and 11/6300 for the trend case respectively. The major shortcoming of Hadri's panel unit root test is that the test statistic does not remain valid under small N and moderate T.