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between health care expenditure and GDP? An
analysis using regime shifts**

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Abstract

This paper investigates the impact of policy shifts on disaggregated health expenditure-GDP relationship for Australia and the USA. In contrast to previous studies the disaggregation is at the level of type of service delivered and not at the level of source of expenditure. Our results show that the subcomponents of health expenditure exhibit different patterns of behaviour at both cointegration and unit root stages once policy shifts or structural breaks, such as the introduction of a publically funded medicare policy in the USA, are allowed in the empirical analysis. When the possibility of structural break is allowed we find a significant long run relationship between subcomponents of aggregate health expenditure and GDP that is not found when no break is considered. The underlying reasons for the occurrence of breaks and policy lessons are discussed subsequently.

1 Introduction

The relationship between aggregate health care expenditure and national income (GDP) has received growing attention in the literature (Newhouse, 1977; Hansen and King, 1996; Getzen, 2000; Clemente et al., 2004; Okunade and Murthy, 2002). However, studies on the relationship between income and the *subcomponents* of health expenditure are almost non-existent. We believe that an analysis of the subcomponents will provide important information so that policy makers can make better informed and more efficient decision. This paper analyses the long-run relationship between the subcomponents of health expenditure and GDP for Australia and the USA. These two countries are chosen for the analysis for two reasons: their disaggregated expenditures are directly comparable as they use the same definition to categorise the subcomponents of aggregate health expenditure; and such data for other OECD countries are either unavailable or available for a very short time span which makes it unsuitable for time series analysis. We argue that in addition to the aggregate health expenditure-GDP relationship, it is critical to analyse the linkages between the subcomponents of health expenditure and GDP. We argue that such an analysis is particularly interesting for two reasons.

Firstly, it is likely the subcomponents which are subject to individual policies and reforms (e.g. pharmaceutical benefit scheme) behave differently from the aggregate health expenditure. For example, dental care is not covered by government-funded policy in most developed countries and thus dental expenditures are either incurred directly by individuals or indirectly through ancillary benefits provided through private health insurance. Thus, dental expenditure might behave differently from the other government-regulated subcomponents of health expenditure.

Secondly, governments face different incentives for allocating funds to various areas of health care which may alter the long-run relationship between national income and the sub-

components of health expenditure. Thus, the public-funded subcomponents of aggregate health expenditure could be subjected to different policy regimes. For example, around 80% of prescriptions dispensed in Australia are subsidised under the publically funded Pharmaceutical Benefits Scheme (PBS). The criteria for subsidising a drug under PBS include the medical conditions for which the drug has been approved for use in Australia, its clinical effectiveness, safety and cost-effectiveness, which is different from the criteria for funding hospitals. Thus, pharmaceutical expenditures and the other subcomponents of aggregate health expenditure may follow different behavioural patterns.

The recent literature on health expenditure has largely focussed on the non-linear time series properties of health expenditure or the presence of non-linearities in its relationship with GDP (Jewell et al., 2003a; MacDonald and Hopkins, 2002; Clemente et al., 2004; Narayan, 2006). However, all of the studies with the exception of Clemente et al. (2004) have used aggregate health expenditure for empirical analysis. Clemente et al. (2004) have specified aggregate health functions separately for private and government expenditures. In contrast, this paper analyses the linkages between income and the subcomponents of total health expenditure (viz. hospital, medical, pharmaceutical and dental expenditures). Thus, the disaggregation is done by type of service delivered and not by source of expenditure. In addition, since health policy shifts can potentially have important implications for health expenditure modelling, the econometric analysis in this paper allows for the impact of policy/regime shifts or structural breaks at both unit root and cointegration stages.

Both the Australian and the USA health expenditures and their subcomponents have been susceptible to structural breaks in response to significant government and other policy changes in their health care sector over the last five decades. For example, a comprehensive government-funded medicare policy was implemented in 1983-84 in Australia and a set of private health insurance incentives schemes were introduced in 2000-01. Similarly in

dentistry, a commonwealth national dental program was introduced in 1993-94 but was subsequently discontinued in 1996. In the USA, major health policy changes involved the introduction of the Medicaid and Medicare legislation in 1965 and a prospective payment system for hospital funding in the early 1980s. In this paper, break dates are assumed to be endogenous (i.e. data dependent) making use of recent advances on endogenous structural breaks in unit root and cointegration literature.

The rest of the paper is organised as follows. Section 2 tests the unit root properties of health expenditure and GDP series firstly using the standard approach and secondly by allowing for endogenous structural breaks as suggested by Zivot and Andrews (1992). Section 3 uses both the standard cointegration approach and residual based methods incorporating regime shifts (Gregory and Hansen, 1996) to detect any long-run relationship between GDP and the non-stationary components of health expenditure. Section 4 concludes.

2 Unit roots and regime shifts

The data used in this paper comprise real per capita annual aggregate health expenditure and four of its subcomponents (i.e. hospital, medical services, pharmaceutical and dental) and GDP. The data for Australia are taken from the Australian Institute of Health and Welfare (AIHW, 2005) and, that of USA, from the Centers for Medicare and Medicaid Services (CMS, 2007). The data series are in natural logarithm for the years 1971-2004 for Australia and 1960-2005 for the USA. Initially, we apply the standard Augmented Dickey Fuller (ADF) stationarity test. The null hypothesis of unit root cannot be rejected at conventional levels of significance. Thus, all series are found to be non-stationary (I(1)).

We next test for structural breaks. While the test can be extended to multiple structural breaks, we restrict our analysis to one structural break. This is done for several reasons: following a recent revision of the OECD health expenditure series, consistent

data are only available from 1970 onwards making the sample insufficiently large for multiple break analysis. Almost all of the previous studies, prior to the revision of the data used 1960-1997 sample with a sample size of 38 for time series analysis. Most of these studies incorporated a single structural break and multiple structural breaks were allowed only in panel unit root tests. For example, the most recent contribution to the relevant literature by Narayan and Narayan (2008) has used one structural break Zivot and Andrews test (Zivot and Andrews, 1992) to test for the unit root properties of aggregate health expenditure. The multiple breaks tests suggested by Lumsdaine and Papell (1997) and Lee and Strazicich (2003) require a large sample size (to retain power of tests and asymptotic properties) which is not possible with the health data extensively used in the literature. It is not possible to explore multiple structural break in a panel setup in our case as unfortunately the data on the subcomponents of health expenditure are not yet available for many OECD countries. Also it should be noted that the focus of this study is to highlight the importance of regime shifts on the unit root properties and co-movement of health expenditures and not to explore number of regime shifts.

A structural break is said to have taken place when a change is observed in the regression parameters of the model. Several studies (For example Zivot and Andrews (1992) and Lumsdaine and Papell (1997)) have reported instances where a series classified as non-stationary (i.e., $I(1)$) in the absence of a structural break hypothesis is actually trend-stationary once structural breaks are accounted for. Thus, the conventional unit root tests erroneously fail to reject the null of unit root. Structural breaks result from one or a few events that significantly affect a variable or set of variables. Such breaks can lead to a permanent shift in the level or slope (or both) of the series but the basic nature of the series remains unchanged. With such events or shocks accounted for, the series can be trend-stationary but with a structural break. Nelson and Plosser (1982) initiated the contemporary debate on structural breaks. They opined that shocks have a permanent effect

on the long-run level of macroeconomic series. This challenged the then prevailing view that shocks subside after a period of time and in the long run any series returns to its trend path. Perron (1989) admitted the possibility of (exogenous) structural breaks in the Nelson and Plosser data series and claimed that the conventional unit root test could fail to reject the unit root hypothesis of non-stationarity even for series known to be trend-stationary with structural break. He proposed a modified version of the conventional ADF test to rectify this.

Zivot and Andrews (1992) criticized Perron's assumption of an exogenous date of structural break and allowed the date of the structural break to be endogenously determined within the model. This reversed some of the results of Perron: They failed to reject the unit root null for four of the series which Perron had classified as stationary. The basic specification of the Zivots and Andrews model for any time series H_t is:

$$\Delta H_t = \alpha + \beta t + \gamma DI_t + \omega DS_t + \mu H_{t-1} + \sum_{i=1}^k c_i \Delta H_{t-i} + \epsilon_t \quad (1)$$

for $t = 1, \dots, T$; where $c(L)$ is a lag polynomial of known order k and $1 - c(L)L$ has all its roots outside the unit circle. DI_t is the indicator dummy variable for a mean shift occurring at time SB and DS_t is the corresponding trend shift variable such that $DI_t = 1$ for $t > SB$. $DS_t = t - SB$ if $t < SB$. The k extra regressors are included to address the problem of autocorrelation, i.e., the temporal dependence in the error terms. A test of the unit root hypothesis has the null $\mu = 0$. The alternative hypothesis is that the series is $I(0)$ with one structural break.

In these specifications, the choice of the lag length k is crucial. Hall (1994) suggests that for moderate to large samples a general-to-specific approach performs better than standard information criteria such as those due to Hannan, Quinn and Akaike and Schwarz. We therefore use the general to specific approach adopted by Perron (1989). The date of the structural break (SB) is allowed to vary between $t = 2$ to $T - 1$ where T is the time period.

We assign a dummy variable for each value of $t \in [2, T-1]$. Hence we get $T-2$ combinations of the data set. The Zivot and Andrews procedure for deciding the date of the structural break chooses that period as the break point which supports the alternative hypothesis the most, i.e., supports the null hypothesis the least. To do this we run a sequential OLS procedure and test for the significance of the coefficient of H_{t-1} (i.e., whether $\mu = 0$). We thus get $T - 2$ t-statistics along with the corresponding coefficients. To decide on SB , the date of the structural break, we choose the minimum value of one-sided (left-tailed) t-statistic calculated. The date corresponding to this minimum t-statistic is chosen as the date of the structural break. Thus, $t(SB) = \min_{T_i} t_\mu(T_i)$. The results for stationary tests incorporating structural breaks are reported in Table 1. The significance of the t-statistics is tested using finite sample critical values (Narayan, 2005; Jha and Sharma, 2004) derived by Monte Carlo simulations as recommended in Zivot and Andrews (1992). We have also tested the unit root properties using the test proposed by Lee and Strazicich (2004) which is more robust for determining structural break dates. The results are qualitatively the same and are skipped here for brevity.

—*Insert Table 1 about here*—

The results for Australia reveal that medical expenditure is trend-stationary with a structural break. The date of break is year 2000 which coincides with a period of major policy reforms of the Australian health care system. Note that in the period 1999 to 2001, several private health insurance incentives schemes were introduced to encourage membership of private health insurance.

In contrast, aggregate health expenditure and its other subcomponents show different patterns of behaviour and are non-stationary. We thus conclude that the effects of structural breaks or regime shifts on the subcomponents can be different from that on

aggregate health expenditure. As for the USA, the stationarity test results reveal that aggregate health expenditure and its subcomponents are all non stationary.

The main implication of the stationary tests is that the structural breaks did not have a permanent effect on the behaviour of one of the subcomponents of health expenditure. The evidence on the stationarity of health expenditures series is mixed in the current literature. For example, our result of non stationarity of US health expenditure is in agreement with the finding by Okunade and Murthy (2002). On the other hand Jewell et al. (2003b) and Carrion-i Silvestre (2005) find US health expenditure to be stationary with breaks. It should be noted that our results are not directly comparable with those of the above studies which have used panel unit root tests and a different time period (1960 to 1997). Our results also have implications for the econometric modelling of the health-GDP relationship. The relationship between the trend-stationary component of health expenditure and GDP cannot be estimated using the cointegration approach as this might lead to misleading results. The cointegration approach is most appropriate for non-stationary series. We next analyse the impact of regime shifts on the long-run relationship between the non-stationary components of health expenditure and GDP.

3 Cointegration with endogenous regime shift

The above analysis established that the health expenditure and most of its subcomponents are non stationary. There is an important literature (Chow, 1960; Quandt, 1960) which emphasises the need to allow for changes in the values of estimated parameters over time, in time series regressions. However, these studies do not consider models with non-stationary regressors. Gregory and Hansen (1996) suggest residual-based tests for cointegration allowing for regime shifts (See Esteve and Sanchis-Llopis (2005); Jha and Sharma (2004) for other applications of this method). These tests reject the null hypothesis of no cointegra-

tion if there is cointegration with regime shifts.

These residual-based tests incorporate regime shifts in the trend as well as the slope of the standard cointegration model. The three specifications to model relationship between health expenditure (H_t) and GDP (Y_t) are:

Standard Cointegration:

$$H_t = \mu + \alpha^\top Y_t + e_t, \quad t = 1, \dots, n \quad (2)$$

Level shift (C):

$$H_t = \mu_1 + \mu_2 \varphi_{t\tau} + \alpha^\top Y_t + e_t, \quad t = 1, \dots, n \quad (3)$$

Level shift with trend (C/T):

$$H_t = \mu_1 + \mu_2 \varphi_{t\tau} + \beta t + \alpha^\top Y_t + e_t, \quad t = 1, \dots, n \quad (4)$$

Level and slope shift (C/S):

$$H_t = \mu_1 + \mu_2 \varphi_{t\tau} + \alpha_1^\top Y_t + \alpha_2^\top Y_t \varphi_{t\tau} + e_t, \quad t = 1, \dots, n \quad (5)$$

Standard cointegration is tested using the Said and Dickey (1984) *ADF* test. Gregory and Hansen (1996) suggest three tests to detect cointegration in the possible presence of regime shifts. The first two are the modified Phillips (1987) test statistics ($Z_t(\tau)$, $Z_\alpha(\tau)$) and the third one is the date-dependent *ADF* statistic (ADF^*). We focus on the ADF^* test here as Gregory and Hansen (1996) recommend comparison of *ADF* and ADF^* to detect the importance of structural break in the long-run relationship. An ADF^* statistic is computed for each possible structural change and the smallest value (largest negative value)¹ is taken across all the break points. Thus:

$$ADF^* = \inf_{t \in \tau} ADF(\tau) \quad (6)$$

¹The smaller is the test statistic the stronger is the evidence against the null hypothesis

The significant value of the test statistic implies rejection of the null of no cointegration against the alternative of cointegration with regime shift. In addition, rejection by either the standard ADF or ADF^* implies that there is some long-run relationship between the two series. The test for multiple breaks in a cointegration setup are only available in a panel framework which could not be used due to lack of availability of large panel data. The approach used in our analysis can however capture up to three policy shifts depending on the model specification which is sufficient for the sample size used here. Results for all three types of models for ADF and ADF^* are presented in Table 2 for the USA and Table 3 for Australia.

—*Insert Table 2, 3 about here*—

For the USA the standard ADF test (with or without level shift) indicates a long run relationship between aggregate health expenditure and GDP but not for the subcomponents of health expenditure. This result of a long run relationship between aggregate health expenditure and GDP for US is consistent with the previous literature. For example, Okunade and Murthy (2002) find a stable long run relationship between aggregate health expenditure and GDP once accounting for the technological change.

However, the ADF^* rejects the null of no cointegration against the alternative of the presence of cointegration with unknown regime shifts for aggregate health expenditure, dental and pharmaceutical expenditures. According to Gregory and Hansen (1996), “*if the standard ADF statistic does not reject the null, but the ADF* does, this implies that structural change in the cointegrating vector may be important*”. Thus, the ADF^* test reveals that if regime shift is allowed for, there is evidence of a long-run relationship between GDP and dental expenditure series. The data dependent dates of regime shifts are 1966 (one year after the Medicaid and Medicare legislation was introduced) and 1983 (the period when dental use slowed from their historical growth rates almost to a standstill

(For example, see Brown et al. (1994))). Similarly, a long run relationship is found between GDP and pharmaceutical expenditure with date of regime shift as 1980. On the other hand, the existence of a long-run relationship between aggregate health expenditure and GDP is supported by both ADF and ADF^* tests and thus a definitive conclusion about the importance of a break point on the relationship between aggregate health expenditure and GDP cannot be made on the basis of this evidence alone.

Similarly for Australia, the results reveal that dental expenditures have a long-run relationship once regime shifts are allowed for in the analysis. The dates of breaks are 1979 (Oil price shock and recession in the economy) and 1993 (introduction of the Commonwealth dental program). Similarly, incorporation of regime shifts reveal a long-run relationship between GDP and pharmaceutical expenditure with breakpoint 1983 (for level and slope specification) and 1978 (for level and level with trend specifications). Note that in 1978 a National Health Act was introduced to allow dentists to prescribe a limited range of antibiotics, antibacterial and antifungal drugs as pharmaceutical benefits. In 1983 a concessional beneficiary category was created to assist the disadvantaged where low-income earners and the unemployed (now concession cardholders) would pay a concessional amount for listed pharmaceutical.

Table 3 shows that both ADF and ADF^* tests support the existence of a long-run relationship of aggregate health expenditure and hospital expenditure with GDP. Thus a definitive conclusion about the importance of a break point on the long run relationships cannot be made on the basis of this evidence alone. The results from the cointegration analysis highlight the importance of regime shifts in detecting the long-run relationship between the subcomponents of health expenditure and GDP. For example the standard cointegration tests reveal the absence of a long-run relationship between dental expenditure and GDP which is counterintuitive as both in the USA and Australia dental services are

paid directly by consumers and are expected to be sensitive to economic cycles and hence GDP. The results are reversed once regime shifts are allowed for in the empirical model.

4 Conclusion

This paper investigates the impact of policy shifts on disaggregated health expenditure-GDP relationship for Australia and the USA. In contrast to previous studies the disaggregation has been done at the level of type of service delivered and not at the level of source of expenditures. Our results show that the subcomponents of health expenditure exhibit different patterns of behaviour at both cointegration and unit root stages once regime shifts are allowed for in the empirical analysis. Thus, our results suggest that the use of standard cointegration and unit root tests (which do not allow for regime shifts) on aggregate health expenditure might result in misleading findings. The results further confirm that for policy purposes the analysis of the health expenditure and GDP linkage should be undertaken separately by disaggregating the aggregate health expenditure by type of service delivered.

What are the policy lessons derived from our study? The impact of shifts in health policies on health expenditure GDP relationship is noteworthy for both Australia and the USA. Dental expenditure is not generally covered by any public insurance scheme neither in Australia nor in the USA. Yet, our results show that the introduction of a public policy such as the Medicaid and Medicare legislation in the USA (which is intended to affect public expenditure) led to both level and slope shift in the privately funded ² dental expenditure-GDP long run relationship. Such an evidence indicates a substitution effect between public and private expenditures. Similarly, for Australia the recession of late 70s was responsible for a level shift in the dental-GDP relationship. This suggests that dental expenditures are pro-cyclical. Such a result indicates the importance of recessionary policies on private

²Total CMS spending on dental care including Medicaid is less than 6 percent of the total to date.

health expenditures.

The results also indicate the effectiveness of policy shifts: the introduction of the Commonwealth dental program significantly affected not only the level but also the trend of the long run dental-GDP relationship. Similarly, the introduction of some new pharmaceutical policies also had significant impacts on the long-run relationship between pharmaceutical expenditure and GDP. For instance, the policy of allowing dentists to prescribe a limited range of drugs as pharmaceutical benefits in Australia significantly affected the level and trend of long run relationship between pharmaceutical expenditure and GDP whereas the policy of creating a concessional beneficiary category to assist the disadvantaged for listed pharmaceutical significantly affected the slope and trend of the long run relationship.

This paper emphasizes the importance of examining health expenditure at a disaggregated level. We believe that the subcomponents of health expenditure are likely to behave differently and examining health expenditure in an aggregated form fails to reveal such differences. Thus policymakers should focus on different subcomponents of health expenditures to formulate more effective policies. Moreover, any empirical analysis to explore long run relationships should allow for unknown breaks. When the possibility of structural breaks is allowed we find a significant long run relationship between subcomponents of aggregate health expenditure and GDP that is not found when break is not considered. Lastly, the evidence of public private substitution at a macro level is a significant finding and could be further explored by using micro level individual data.

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Table 1: Results: Stationarity with Single Endogenous Break

Series	Min. t-stats (μ)	Lags (k)	Date of Break	Comments
Australia (1971-2004)				
Total Health Exp.	-6.81	3	-	Non-stationary
Hospital Expenditure	-4.64	5	-	Non-stationary
Medical Expenditure	-8.47***	3	2000	Trend-stationary
Pharmaceutical Exp.	-4.45	7	-	Non-stationary
Dental Expenditure	-4.08	0	-	Non-stationary
GDP	-4.01	1	-	Non-stationary
USA (1960-2005)				
Total Health Exp.	-5.45	8	-	Non-stationary
Hospital Expenditure	-3.20	1	-	Non-stationary
Medical Expenditure	-4.60	1	-	Non-stationary
Pharmaceutical Exp.	-5.62	3	-	Non-stationary
Dental Expenditure	-4.47	1	-	Non-stationary
GDP	-4.67	1	-	Non-stationary

Finite Sample Critical values for US Data: -8.36 (1%), -7.50 (5%), -7.15 (10%)

Finite Sample Critical values for Australian Data: -8.43 (1%), -7.55 (5%), -7.25 (10%)

***:significant at 1% level.

- : Not Applicable

Table 2: Results: Cointegration with Regime Shifts: USA

Tests	Dental - GDP		Medical - GDP		Hospital - GDP	
	Test Stat	Breakpoint	Test Stat	Breakpoint	Test Stat	Breakpoint
<i>ADF*</i>						
C	-4.63**	1983	-4.04	-	-3.79	-
C/T	-3.81	-	-3.55	-	-3.99	-
C/S	-5.00**	1966	-4.64	-	-4.45	-
ADF						
C	-3.11	-	-2.38	-	-1.52	-
C/T	-2.32	-	-1.88	-	-1.36	-
Tests	Health Exp. - GDP		Pharmaceutical Exp. - GDP			
	Test Stat	Breakpoint	Test Stat	Breakpoint		
<i>ADF*</i>						
C	-4.62*	1995	-2.74	-		
C/T	-4.28	-	-3.27		-	
C/S	-4.63	-	-4.69*	1980		
ADF						
C	-3.90**	-	-1.70	-		
C/T	-2.41	-	-3.19	-		

***:significant at 1% level.

**:significant at 5% level.

*:significant at 10% level.

The significance levels of *ADF** are calculated by using asymptotic critical values (for $m = 1$) reported in Table 1 of Gregory and Hansen (1996)

- : Not Applicable