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Recent Trends in Fetal and Infant Outcomes Following Post-term Pregnancies

*Shi Wu Wen, K S Joseph, Michael S Kramer, Kitaw Demissie, Lawrence Oppenheimer, Robert Liston and Alexander Allen for the Fetal and Infant Mortality Study Group, Canadian Perinatal Surveillance System**

Abstract

All births and infant deaths in1985–87 and 1992–94 in Canada, except in Ontario and Newfoundland, were analyzed to assess the potential impact of the recent increased use of elective labour induction for post-term pregnancies. Probabilistic linkage was carried out of infant death records (Canadian Mortality Database) and respective birth registrations (Canadian Birth Database) for the periods 1985–87 and 1992–94. The combined fetal and infant mortality declined by 20–30% between 1985–87 and 1992–94 at each gestational week beginning at 37 weeks, with no increased reduction among post-term pregnancies. Asphyxia-related fetal and infant deaths, the most likely cause of death being preventable by labour induction for post-term pregnancies, did not decrease among post-term pregnancies. On the contrary, a substantial decrease of asphyxia-related deaths was observed at 37 and 38 weeks over the same periods of time. Because fetal and infant deaths are rare events and because the number of pregnancies passing 42 weeks of gestation decreased dramatically during 1992–94, statistically unstable results may be inevitable in the comparison of mortality in this group of pregnancies.

Key words: *labour induction; mortality; post-term pregnancy*

Introduction

Post-term pregnancies are pregnancies that reach at least 42 weeks of gestation.[1] Perinatal mortality and the occurrence of various obstetric complications have been found to be higher in post-term than in term pregnancies.[2,3] Two general management approaches have been developed to reduce the risk of these adverse outcomes: elective labour induction when the pregnancy reaches 41 or 42 weeks of gestation, or expectant management with frequent fetal monitoring and selective labour induction. The two approaches remain controversial.[4–9]

Proponents of elective labour induction cite evidence from randomized controlled trials showing that labour induction is associated with reduced perinatal mortality.[4–6] Concerns have been raised, however, that results from tightly controlled trials may not be applicable in routine practice and that widespread implementation of routine induction for post-term pregnancies may lead to increased rates of cesarean section and other obstetric interventions.[7–9] Despite these controversies, the use of elective labour induction for post-term pregnancies has increased dramatically in Canada since the early 1990s, mostly at 41weeks of gestation.[6] During this period of time, to our knowledge, the only apparent difference in obstetric care between post-term and term pregnancies in Canada was a tendency to induce labour electively for the former, but not for the latter.

Author References

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A recent observational study suggests that the increased use of elective labour induction for post-term pregnancies has contributed to the decline in fetal deaths among these pregnancies in Canada.[6] However, this study did not assess infant mortality or causes of fetal death. We hypothesized that if routine, elective labour induction in recent years has had an important impact on the fetal and infant outcomes of post-term pregnancies in the form of the reduction of fetal and infant mortality, of asphyxia-related deaths compared to other causes of death, and of perinatal compared to post-neonatal death, it would be more evident in post-term pregnancies than in term pregnancies. We carried out an epidemiologic study using linked birth and infant death files to test these hypotheses.

Methods

We used data for live births and stillbirths from Statistics Canada's Canadian Birth Database[10] for the years 1985–1994 and data for fetal and infant deaths from the Canadian Mortality Database for the years 1985–1995. Fetal death is defined as stillbirth with birth weight ≥ 500 g or gestational age ≥ 20 weeks; neonatal death is defined as a live birth of an infant that died before the 28th full day of life; post-neonatal death is defined as a live birth of an infant that died between the 29th and the 364th full day of life; and infant death is defined as a live birth of an infant that died before the 364th full day of life. A probabilistic linkage was carried out using previously validated methods to link infant death records with respective birth registrations.[11,12] Uncertain linkages were resolved after a manual examination of the relevant birth and death registration documents.

Ontario births were excluded from the analysis because of documented problems with data quality.[13] Newfoundland births were also excluded from the analysis of time trends, because data from this province were not available before 1991. Information in the linked files of live births and infant deaths was subjected to internal data quality checks, including procedures to exclude duplicate records.

The linked birth and death information enabled the creation of birth cohorts with follow-up information on mortality in the first year after birth. Thus, infants born in 1985 were followed through 1986 to calculate infant mortality rates. Similarly, while the last birth cohort constructed was of live births in 1994, follow-up for infant death among newborns in this cohort extended through 1995.

For the current study, only term and post-term births (i.e., those of 37 or more weeks of gestation) were included in the analysis. Because fetal and infant mortality decreased continuously during the 10-year period covered by the study and differences between successive years were small, we combined the data for 1985–1987 births and 1992–1994 births to enhance the statistical stability of the estimates. We calculated rates of fetal death, neonatal death, post-neonatal death, and overall fetal and infant death rates in each gestational week for the two study periods. Relative risks and 95% confidence intervals were used to compare mortality rates between the two study periods, with 1985–87 serving as the reference.

We further analyzed fetal and infant mortality caused by asphyxia, on which labour induction probably has the largest impact. Only one underlying cause of death is recorded in Statistics Canada's Canadian Mortality Database, which is coded using the International Classification of Diseases, 9th Revision (ICD-9) classification system. We used a classification system adopted by the International Collaborative Effort on Perinatal and Infant Mortality[14] to group codes with clinical conditions that may directly or indirectly lead to asphyxia-related death. These conditions include maternal death (ICD-9 761.6), malpresentation before labour (ICD-9 761.7), placenta previa or other placental abnormalities (ICD-9 762.0-762.2), prolapsed cord or other unspecified conditions of umbilical cord (ICD-9 762.4, 762.5, 762.6), breech delivery and extraction (ICD-9 763.0), disorders relating to long gestation and high birth weight (ICD-9 766), birth trauma (ICD-9 767), intrauterine hypoxia and birth asphyxia (ICD-9 768), meconium aspiration syndrome (ICD-9 770.1), subarachnoid hemorrhage (ICD-9 772.2), convulsions in newborn (ICD-9 779.0) and coma or other abnormal cerebral signs (ICD-9 779.2).

We considered the potential impact of the reclassification of gestational age among post-term pregnancies, caused by more frequent labour induction at 41 or more weeks of gestation,[6] on gestational age-specific mortality, and then carried out a parallel analysis combining all births at 41 or more weeks and compared the results with the main analysis using finer groupings of gestational age (37, 38, 39, 40, 41, 42 and ≥ 43 weeks of gestation).

Results

The number of births and deaths are presented in Tables 1 and 2. Compared with 1985–87, births at 37, 38, 39 and 41 weeks of gestation increased as a proportion of total births in 1992–94, but decreased at 40 and (especially) 42 and ≥43 weeks (Table 1)

The results of the comparison of fetal and infant mortality rates between 1985–87 and 1992–94 are presented in Table 3. The fetal death rate at 37 weeks decreased by 28% from 1985–87 to 1992–94. The reduction in fetal death rates increased sequentially from 40 weeks onwards, being 12%, 15%, 20% and 35% among those at 40, 41, 42 and ≥ 43 weeks respectively. On the other hand, the decrease in neonatal mortality was greatest at 40 weeks (39% decrease in 1992–94 versus 1985–87; see Table 3). Compared with births at 40 weeks, the decreases were smaller at 37, 38, 39 and 41 weeks and there was no statistically significant reduction at 42 weeks (relative risk [RR] 0.94, 95% confidence interval [CI] 0.64-1.37). At ≥ 43 weeks, there was even a statistically nonsignificant increase. Post-neonatal

TABLE 1
Gestational age distribution of term and post-term births (combining stillbirths and live births), Canada excluding Ontario and Newfoundland, 1985–87 and 1992–94

Gestational age (weeks)	1985–87		1992–94		Percent change
	Number	Percent	Number	Percent	
37	33,143	4.76	40,444	5.69	+21.0
38	88,137	12.65	102,087	14.36	+14.8
39	136,154	19.54	153,215	21.56	+11.6
40	259,698	37.27	230,567	32.44	–12.0
41	95,505	13.71	104,147	14.65	+8.2
42	34,549	4.96	24,161	3.40	–30.5
≥43	2,878	0.41	956	0.13	–65.9
Total	**650,064**		**655,577**		

TABLE 2
Number (rate per 1,000) of fetal death, neonatal death, post-neonatal death, total fetal and infant death, and fetal and infant death for asphyxia-related conditions by gestational age in term and post-term births, Canada excluding Ontario and Newfoundland, 1985–87 and 1992–94*

Gestational age	Fetal death		Neonatal death		Post-neonatal death		Fetal and infant death		Asphyxia-related fetal and infant death	
	1985–87	1992–94	1985–87	1992–94	1985–87	1992–94	1985–87	1992–94	1985–87	1992–94
37	232 (7.0)	205 (5.1)	134 (4.1)	139 (3.5)	129 (3.9)	121 (3.0)	495 (14.9)	465 (11.5)	143 (4.3)	124 (3.1)
38	297 (3.4)	255 (2.5)	204 (2.3)	174 (1.7)	270 (3.1)	259 (2.5)	771 (8.8)	688 (6.7)	184 (2.1)	156 (1.5)
39	239 (1.8)	239 (1.6)	213 (1.6)	167 (1.1)	296 (2.2)	256 (1.7)	748 (5.5)	662 (4.3)	167 (1.3)	160 (1.0)
40	360 (1.4)	280 (1.2)	349 (1.4)	189 (0.8)	515 (2.0)	353 (1.5)	1,224 (4.7)	822 (3.6)	257 (1.0)	200 (0.9)
41	157 (1.6)	146 (1.4)	141 (1.5)	98 (0.9)	200 (2.1)	130 (1.3)	498 (5.2)	374 (3.6)	105 (1.1)	100 (1.0)
42	84 (2.4)	47 (2.0)	67 (1.9)	44 (1.8)	81 (2.4)	29 (1.2)	232 (6.7)	120 (5.0)	57 (1.7)	37 (1.5)
≥43	23 (8.0)	5 (5.2)	7 (2.5)	3 (3.2)	8 (2.8)	0 (0.0)	38 (13.2)	8 (8.4)	13 (4.5)	5 (5.2)
≥41	264 (2.0)	198 (1.5)	215 (1.6)	145 (1.1)	289 (2.2)	159 (1.2)	768 (5.8)	502 (3.9)	175 (1.3)	142 (1.1)
Total	**1,392 (2.1)**	**1,177 (1.8)**	**1,115 (1.7)**	**814 (1.2)**	**1,499 (2.3)**	**1,148 (1.8)**	**4,006 (6.2)**	**3,139 (4.8)**	**926 (1.4)**	**782 (1.2)**

*Fetal death is defined as stillbirth with birth weight ≥ 500 g or gestational age ≥ 20 weeks; neonatal death is defined as a live birth of an infant that died prior to the 28th full day of life; post neonatal death is defined as a live birth of an infant that died between the 29th and the 364th full day of life; infant death is defined as a live birth of an infant that died prior to the 364th full day of life.

mortality declined at every week of gestation in 1992–94 compared to 1985–87; this decrease was most evident at 41 or more weeks of gestation.

The combined fetal and infant mortality rate had decreased by 20%–30% for every week of gestation in 1992–94 compared to 1985–87. The reduction at 41 or more weeks of gestation was slightly larger than the reduction at 37–40 weeks, mainly because of a larger reduction in post-neonatal mortality rates at these gestations.

In general, fetal and infant mortality caused by asphyxia decreased in 1992–94 compared to 1985–87. However, the difference was statistically significant only for 37 and 38 weeks of gestation. Results obtained from analysis combining all births at 41 or more weeks (second row from bottom of Table 3) were generally consistent with those using finer categories of post-term pregnancies.

Discussion

Our study showed a moderate increase in the proportion of births at 41 weeks of gestation and substantial decreases in this proportion at 42 weeks and ≥ 43 weeks in 1992–94 compared to 1985–87. These changes are probably partly attributable to better dating of pregnancies with more frequent use of ultrasound early in pregnancy,[15,16] although more frequent labour induction at ≥ 41 weeks has doubtless also played a role. Although we have no direct data on labour induction, a recent study showed a substantial rise in rates of labour induction at 41 weeks from the early 1990s in the majority of Canadian hospitals surveyed.[6] The continued increase in inductions among post-term pregnancies after

TABLE 3
Relative risk (95% confidence interval)* for fetal death (per 1,000 total births), neonatal death (per 1,000 live births), post-neonatal death (per 1,000 survivors 28 days of age), total fetal and infant death (per 1,000 total births), and fetal and infant death (per 1,000 total births) for asphyxia-related conditions by gestational age in term and post-term births, Canada excluding Ontario and Newfoundland

Gestational age	Fetal death	Neonatal death	Post-neonatal death	Fetal and infant death	Asphyxia-related fetal and infant death
37	0.72 (0.60, 0.87)	0.85 (0.67, 1.08)	0.77 (0.60, 0.98)	0.77 (0.67, 0.87)	0.71 (0.55, 0.91)
38	0.74 (0.63, 0.88)	0.74 (0.60, 0.90)	0.83 (0.70, 0.98)	0.77 (0.69, 0.85)	0.73 (0.59, 0.91)
39	0.89 (0.74, 1.06)	0.70 (0.57, 0.85)	0.77 (0.65, 0.91)	0.79 (0.71, 0.87)	0.85 (0.69, 1.06)
40	0.88 (0.75, 1.02)	0.61 (0.51, 0.73)	0.77 (0.67, 0.88)	0.75 (0.69, 0.83)	0.88 (0.73, 1.05)
41	0.85 (0.68, 1.07)	0.64 (0.49, 0.82)	0.60 (0.48, 0.74)	0.69 (0.60, 0.79)	0.87 (0.66, 1.15)
42	0.80 (0.56, 1.14)	0.94 (0.64, 1.37)	0.51 (0.33, 0.78)	0.74 (0.59, 0.92)	0.93 (0.61, 1.40)
≥ 43	0.65 (0.25, 1.72)	1.29 (0.33, 4.97)	0.00 (0.00, 2.03)	0.63 (0.30, 1.35)	1.16 (0.41, 3.24)
≥ 41	0.77 (0.64, 0.93)	0.69 (0.56, 0.85)	0.56 (0.47, 0.68)	0.67 (0.60, 0.75)	0.83 (0.73, 1.04)
Total	**0.84 (0.78, 0.91)**	**0.72 (0.66, 0.79)**	**0.76 (0.70, 0.82)**	**0.78 (0.74, 0.81)**	**0.84 (0.76, 0.92)**

*1992–94 versus 1985–87 rates; fetal death is defined as stillbirth with birth weight ≥ 500 g or gestational age ≥ 20 weeks; neonatal death is defined as a live birth of an infant that died prior to the 28th full day of life; post neonatal death is defined as a live birth of an infant that died between the 29th and the 364th full day of life; infant death is defined as a live birth of an infant that died prior to the 364th full day of life.

1991 has probably been fueled, at least in part, by evidence from randomized controlled trials supporting this practice[4,5] and by the increasing availability of intracervical and vaginal prostaglandin gels to assist with cervical ripening.[6]

During the same period of time, we observed a relatively larger (but statistically nonsignificant) reduction in fetal death among pregnancies at ≥ 41 weeks as compared with those at 40 or 39 weeks. This finding is consistent with a recent Canadian study.[6]

It is notable that the decrease in neonatal mortality at 41 weeks was not greater than at 40 weeks, and that no decrease occurred at 42 or ≥ 43 weeks; in fact at ≥ 43 weeks there was a statistically nonsignificant increase (Table 3). In general, larger reductions in fetal death have been accompanied by smaller reductions in neonatal mortality. For example, the decrease in fetal death was greater at 37 and 38 weeks than at 39 and 40 weeks, whereas the reverse was true for the decrease in neonatal mortality. This pattern of smaller reductions in neonatal mortality than in fetal mortality raises the possibility that more frequent use of medical interventions such as labour induction may have merely postponed some deaths. An alternative explanation for this phenomenon is that labour induction may be beneficial in certain circumstances but harmful in others.

The reduction in post-neonatal mortality in 1992–94 versus 1985–87 was larger at ≥ 41 weeks than at 37 to 40 weeks (Table 3). This finding is in contrast to randomized controlled trials assessing the efficacy of labour induction for post-term pregnancies that have focused on perinatal mortality.[4] It is possible that labour induction may have a larger impact on reducing post-neonatal mortality than it does on perinatal mortality. If this is the case, the assessment of the efficacy of labour induction for post-term pregnancies should be expanded to include the post-neonatal period.

We hypothesized that labour induction for post-term pregnancies would have the largest impact on asphyxia-related fetal and infant mortality and morbidity. The reduction observed in fetal and infant mortality due to asphyxia-related conditions at ≥ 41 weeks, however, was not larger than at 40 weeks (Table 3). During the study period, the clinical definition for certain asphyxia-related conditions such as respiratory distress may have changed. However, our earlier study based on hospital discharge data found that the incidence of coded diagnoses such as respiratory distress and meconium aspiration syndrome was quite stable during the study period.[17] Moreover, since we used a broad and inclusive definition of asphyxia-related deaths, such a shift in clinical definition probably had a limited impact on our study results.

Our population-based results should reflect routine practice better than controlled trials. We realize the limitations inherent in any observational study based on administrative databases. Such data are prone to a certain degree of coding errors,[18] which may be random or may contain systematic biases. The observational study design and the lack of information on induction in the data render any inference about the relation between labour induction and fetal and infant mortality necessarily indirect. Moreover, because fetal and infant deaths are rare events, and because the number of pregnancies passing 42 weeks decreased dramatically during 1992–1994, statistically unstable results were inevitable

in the comparison of mortality in this group of pregnancies. Nonetheless, our findings are biologically plausible.

Trials and guidelines vary widely in terms of gestational age for induction. In the Cochrane systematic review, trials demonstrated an effect of reducing perinatal mortality only for induction conducted after 42 weeks of gestation.[4] The Society of Obstetricians and Gynecologists of Canada initially recommended that women who reach "41–42" weeks of gestation should be offered elective induction,[19] and in practice the tendency has been to induce pregnancies closer to 41 rather than 42 weeks.[6] The most difficult challenge in the management of post-term pregnancy may be related to determining the exact time that a given pregnancy becomes "post-term."[9] It is thus difficult to establish a rigid and arbitrary cut-off point for induction.

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The Duration of Major Depressive Episodes in the Canadian General Population

Scott B Patten

Abstract

The National Population Health Survey (NPHS) has provided a wealth of new data concerning major depression in the Canadian general population. The NPHS included a brief predictor of major depression, and also two questions (only one of which was asked of each subject) concerned with the duration of episodes in the preceding year. A striking finding was that many of the episodes identified were very brief. In this paper the NPHS data were examined from a different perspective in order to derive a complementary perspective on the episode duration data. Data from the 1994/95 and 1996/97 cycles of the NPHS were used in the analysis. The longitudinal data were used to generate approximations of age and gender-specific incidence for members of the population over the age of 12 years. An estimate of prevalence was made from the 1996/97 cross-sectional file. A basic expression relating prevalence to incidence and mean duration of illness was then applied within age and gender categories. Taken together, the incidence and prevalence data from the NPHS suggest a longer duration than was indicated by the NPHS interview duration item. A probable explanation is that the NPHS duration question had an upper limit of 52 weeks, whereas some episodes of major depression last longer than this. Particularly long episodes could have a large impact on mean duration in the population. Nevertheless, these data confirm the heterogenous nature of this condition; many people with the syndrome of major depression may have quite brief episodes.

Key words: *depressive disorder, epidemiology, prognosis, prevalence, incidence*

Introduction

A variety of prevalence estimates for major depression have been reported in various countries.[1–7] Recent estimates of the 12-month period prevalence of major depression in the Canadian population derive from the Canadian National Population Health Survey (NPHS) [8,9] and the Mental Health Supplement of the Ontario Health Survey.[10] Estimates from several additional Canadian surveys have been published.[11,12] Since the NPHS is a longitudinal study, it is possible to approximate incidence using the NPHS data,[13] with the proviso that the follow-up interval for the longitudinal component of the NPHS is two years and the measurement instrument employed identifies episodes occurring only in the 12-month period preceding the interview. The proportion of non-depressed subjects in 1994/95 who were subsequently found to have major depression in 1996/97 may overestimate the annual incidence proportion, since it may include some persons with an onset of major depression in the year following the 1994/95 interview. This would occur if these episodes did not resolve until two weeks or more into the year preceding the 1996/97 interview. These subjects would appear in the numerator of the incidence expression, inflating the estimated incidence proportion.

The NPHS is an ongoing longitudinal community study conducted by Statistics Canada. A national sample consisting of more than 17,000 subjects was interviewed in 1994/95, then re-interviewed in 1996/97 in the longitudinal component. The 1996/97 data collection cycle also included many "buy-ins" from specific provinces, such that the cross-sectional data from the 1996/97 cycle included over 80,000 subjects (a minority of whom were subjects being followed up two years after the initial data collection). The NPHS interview incorporated a short form version of the Composite International Diagnostic Interview (CIDI), called the Short Form for Major Depression.[14] This brief predictive instrument assigns a probability of major depression using a set of questions adapted from the CIDI by Kessler et al.[14] Subjects reporting at least five of nine symptoms

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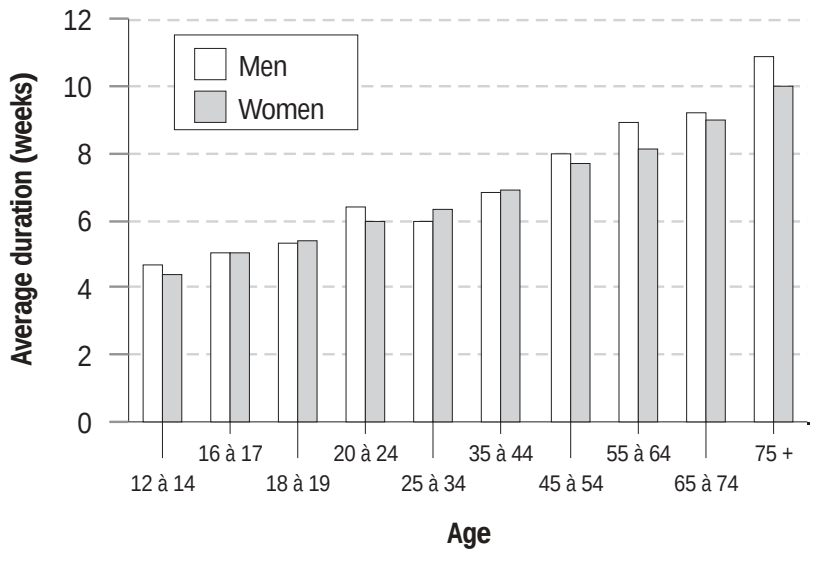
constituting the "A" diagnostic criteria for major depression in DSM-IV, when at least one of these symptoms is depressed mood or loss of interest, are assigned a predictive probability of 90%.

The CIDI short form includes two branched series of questions. Neither branch is followed when a subject does not report a two-week period of depressed mood or a two-week period of loss of interest. Subjects reporting the former symptom follow one branch and subjects reporting the latter symptom (but not the former) follow the other branch. Each branch includes questions about other depressive symptoms requisite to the DSM-IV definition of major depression. Each branch also contains a question about the duration of depression, asking those subjects whose responses indicated an episode of depression to report the total number of weeks in the preceding year that they felt "this way" (the question was preceded by a summarizing statement referring to the occurrence of an episode of depression in the preceding 12 months, and incorporating key phrases for reported symptoms and their duration). Surprisingly brief durations were reported; they are presented in Figure 1.

FIGURE 1
Number of weeks depressed in past 52 weeks*



* Adapted from: Statistics Canada's Internet site <http://www.statcan.ca/english/Pgdb/People/Health/health35.htm>. Date of extraction: November 1, 2000

Potentially, the brief mean duration of depressive episodes has important implications for public health. In clinical settings, the major depressive syndrome is usually regarded as an indication for treatment. Public health efforts, such as the American Depression Screening Day,[15] tend to focus on encouraging people with symptoms suggestive of these disorders to come forward for treatment. Such secondary and tertiary level prevention is an obvious strategy in the absence of feasible approaches to primary prevention. However, both pharmacological and non-pharmacological approaches to treatment generally take at least 4–8 weeks, and the NPHS duration data appear to imply that the average episode would resolve during that time.

The objective of this project was to estimate the mean duration of major depressive episodes using an alternative approach: by integrating the available incidence and prevalence data. This provides an alternative perspective on the duration of these episodes.

Material and Methods

A retrospective cohort analysis was used to estimate age- and gender-specific incidence proportions. Using data collected in 1994/95, all subjects who were under the age of 12 and all subjects with major depression (according to the CIDI-SFMD) were excluded. The remaining subjects were assessed in the 1996/97 data set in order to estimate the proportion of these subjects who had newly developed major depression. Standard errors for these estimates were calculated using a bootstrap method recommended by Statistics Canada. The bootstrap variance estimates were calculated using an SAS[16] macro developed by Statistics Canada for this purpose. All estimates used sampling weights to adjust for unequal selection probabilities due to the complex survey sampling strategies employed by Statistics Canada in this survey.

Prevalence estimates were made from the 1996/97 NPHS data, since the sample size for this data collection cycle was much greater than that of the 1994/95 survey. Again, the data were weighted, this time using a different set of sampling weights produced specifically for the relevant data file. Bootstrap methods were used to calculate confidence intervals for the prevalence estimates.

In order to integrate the incidence data into a broader epidemiological context, an incidence-prevalence model was used. This model represented the relationship between incidence and prevalence in an assumed stable population (and in the absence of migration) as an equality between the inflow of new cases into a prevalence pool and the outflow from the prevalence pool. The outflow includes all terminations of disease, including recovery and death. This model was described by Rothman and Greenland,[17] as follows. In a stable population, the inflow of new cases into the pool of cases (prevalence pool) equals the outflow from the prevalence pool:

$$I(N - P)\Delta t = (1/\overline{D})P\Delta t \qquad \text{(Equation 1)}$$

Where P represents the number of persons in the population with a disorder, I represents the incidence, N the total number of persons in the population, Δt an interval of time and $\overline{D}$ the mean duration of the disorder. This expression may be simplified to:[17]

$$\frac{P}{N-P} = I\overline{D} \qquad \text{(Equation 2)}$$

Here, the ratio on the left side of the equation is the prevalence odds, a parameter that approximates prevalence when a disorder is rare.

For the purpose of this analysis, point prevalence was considered equivalent to one-month period prevalence. One-month period prevalence is often considered equivalent to point prevalence in psychiatric epidemiology because definitions of mental disorders include a requirement not only for the presence of signs and symptoms, but also for the persistence of these signs and symptoms over several weeks. For major depression, the symptoms must persist over at least two weeks. To perform the calculations using months as the time unit, it was necessary to denominate the incidence as a compatible person-month rate by estimating a monthly incidence rate that would be expected to result in the observed incidence proportion over a one-year period. For the purpose of generating this approximation, the "exponential formula" was used:

$$IP \approx 1 - \exp\left(-\sum_{k=1}^{12} I_k \Delta t_k\right) \qquad \text{(Equation 3)}$$

Here, IP is the approximation of the incidence proportion from the longitudinal NPHS data and I_k is a monthly incidence rate that would result in this incidence proportion over a 12-month follow-up interval. The latter estimate is the one suitable for substitution into Equation 2. If episode duration is measured in months, then the I_k rate (which has the units months^{-1}) will result in a dimensionless prevalence odds.

The NPHS survey used a 12-month predictor of major depression prevalence, whereas current prevalence is the parameter of most relevance to the model under development. It was, therefore, necessary to estimate current prevalence by combining the NPHS data with supplementary information from the literature. Fortunately, the literature in this area was found to be strikingly consistent. Kessler et al.,[18] using data from 15–24 year-olds participating in the National Comorbidity Survey (NCS), reported that the ratio of annual to current (30-day) major depression was 12.4% to 5.8%, or approximately 2:1, this ratio being similar in male (9.0% to 3.8%) and female (16.1% to 8.0%) subjects. The overall 12-month period prevalence of major depression in the NCS was 10.3%,[19] compared to the 4.9% one-month period prevalence,[1] a ratio also approximating 2:1. The ratio of annual to current cases in the NCS was similarly comparable among male (7.7% and 3.8%, respectively) and female respondents (12.9% and 4.9%, respectively).

Since the NCS is an American study, the ratio of annual to current major depression was also examined using published data from an earlier survey conducted in Edmonton. This study utilized methods resembling those employed in the Epidemiological Catchment Area (ECA) studies in the United States. Here, an annual prevalence of 4.6% and a one-month prevalence of 2.3% were reported, a 2:1 ratio.[11] The ECA study itself reported a one-year prevalence of (4.2%)[20] and a one-month period prevalence of 2.2%,[4] also closely approximating 2:1. Based on these very consistent findings, the ratio of annual to current major depression in this analysis was taken to be 2:1 for both men and women. Hence, prior to estimation using equation 2, the annual prevalence from the NPHS was multiplied by 0.5 to generate an approximation of point (30-day) prevalence.

Results

The 1994/95 NPHS had a sample size of 17,626. The longitudinal data file included 15,670 subjects providing follow-up data. The current analysis excluded subjects who were under the age of 12 (n = 1908), subjects who had major depression at baseline (n = 781) or who did not provide valid data on the major depression predictor either at the baseline interview, the follow-up interview or both (n = 691). As such, the current analysis of incidence was based on 12,290 subjects. The prevalence estimates were based on 70,538 subjects over the age of 12 in the 1996/97 cross-sectional component of the NPHS. The sample included 73,402 subjects within this age group, but 2,864 (3.9%) who did not provide a valid rating on the CIDI Short Form were excluded.

The age- and gender-specific incidence of major depression, along with 95% confidence intervals, are presented in Table 1. The observed pattern of incidence generally resembled that of prevalence, as depicted in Table 2. Among female subjects, incidence and prevalence

TABLE 1
Age and gender-specific one-year incidence proportions and rates for major depression

		Annual incidence proportion	95% confidence interval	Estimated incidence rate (month^{-1})
Men	Age 12–24	0.029	0.014–0.043	2.45e-03
	Age 25–44	0.033	0.020–0.047	2.80e-03
	Age 45–64	0.018	0.007–0.029	1.51e-03
	Age ≥ 65	0.018	0.007–0.028	1.51e-03
Women	Age 12–24	0.071	0.051–0.091	6.14e-03
	Age 25–44	0.045	0.034–0.057	3.84e-03
	Age 45–64	0.041	0.025–0.057	3.49e-03
	Age ≥ 65	0.013	0.006–0.021	1.09e-03

TABLE 2
Age and gender-specific one-year prevalence proportions and estimated current prevalence rates for major depression

		Annual prevalence proportion (%)	95% confidence interval	Estimated current prevalence proportion*
Men	Age 12–24	2.6%	1.9–3.2	1.30%
	Age 25–44	3.5%	2.8–4.2	1.75%
	Age 45–64	2.6%	2.0–3.2	1.30%
	Age ≥ 65	1.7%	0.9–2.5[M]	0.85%
Women	Age 12–24	6.7%	5.4–8.0	3.35%
	Age 25–44	6.8%	5.9–7.7	3.40%
	Age 45–64	5.0%	4.1–5.8	2.50%
	Age ≥ 65	1.6%	1.0–2.2	0.80%

* 50% of the annual prevalence proportion; see text.

TABLE 3
Estimated duration of major depressive episodes

		Estimated current (30d) prevalence	Estimate incidence rate (month^{-1})	Estimated mean duration of episodes* (months)
Men	Age 12–24	1.30%	2.45e-03	5.4
	Age 25–44	1.75%	2.80e-03	6.4
	Age 45–64	1.30%	1.51e-03	8.7
	Age ≥ 65	0.85%	1.51e-03	5.7
Women	Age 12–24	3.35%	6.14e-03	5.6
	Age 25–44	3.40%	3.84e-03	9.2
	Age 45–64	2.50%	3.49e-03	7.4
	Age ≥ 65	0.80%	1.09e-03	7.4

* Based on equation 2, the mean duration of episodes is calculated as: P / I*(1-P).

tended to be higher in younger women and to decline with advancing age. In males, the incidence rates tended to be slightly higher in middle-aged than in younger subjects. In the subjects aged 75 years and older, the incidence and prevalence increased slightly in both genders.

The prevalence-incidence model embodied in equations 1 and 2 relates incidence and the prevalence proportion (approximately, prevalence) to the mean duration of disease irrespective of whether the disease is terminated by recovery or death. Completing these calculations for each age and gender group yielded the data presented in Table 3. There is no evidence of gender differences in model-based estimates of disease duration, nor is there a pronounced trend towards such differences in association with age. However, the duration of the episodes, as predicted by the incidence-prevalence model, appears to be more brief in the youngest age group.

Discussion

The development of this model incorporated a variety of assumptions. Some assumptions are involved in the estimation of incidence rates using an exponential equation. The use of this equation involves assumptions that the population is closed, that there are no competing risks (often, the equation is used to estimate mortality rates) and that the number of events is small relative to the number at risk.[17] Another assumption was that the 12-month incidence proportion was being measured by the CIDI Short Form. The CIDI Short Form questions are designed to cover a 12-month period, but since the population at risk was identified as those not having major depression during the two previous years it is possible that some of the episodes had their onset more than one year prior to the interview. There may be additional measurement concerns related to the CIDI Short Form. This instrument does not include many of the "clinical significance" probes that are contained in the full CIDI and may, therefore, be less specific in its measurement properties. Finally, the relationship between annual and current prevalence had to rely on data from the literature. Since none of these assumptions can be definitely shown to hold true, the model presented here should be regarded as a heuristic one. It provides a description of the relationship between major depression incidence and prevalence in Canada using the best available data. Furthermore, for a recurrent condition such as major depression, differences in the approach to modelling may be fruitful. For example, the "lifetime sick day proportion"[21] has been proposed as a means of modelling the relationship between incidence and prevalence in episodic conditions accounting both for episode duration and number of episodes. However, such models will also generally be subject to various approximations and assumptions.[21]

Despite these provisos, the model presented here appears to provide a description of incidence-prevalence relationships that are consistent with other available data. The NPHS incidence rates are approximately consistent with those reported elsewhere in the literature. In making such comparisons, it should be emphasized that the NPHS measures major depressive episodes, not disorders, so that the incidence rates are higher than those studies evaluating the first occurrence of depressive episodes (these first episodes being considered the first onset of an episodic depressive disorder in some studies). One recent German study reported a 20-month incidence of major depression in an adolescent sample,[22] 3.7% in male and 7.5% in female respondents. These authors estimated that the 12-month incidence in their sample would have been approximately 4.3%. Another prospective study of high school-aged adolescents reported annual incidence rates for major depression of 10.4% for female and 4.8% for male subjects.[23] One study of the incidence of major

depression in a very elderly (mean age 85) sample reported an annual incidence of 1.4%.[24] All of these findings are very consistent with the data presented here. The six-month follow-up of the Epidemiological Catchment Area survey in New Haven reported a 4.3% six-month incidence.[25] As a six-month incidence estimate (and from a study of older subjects), the New Haven figure is higher than that of the German study, which is a surprising result. The New Haven estimate may, however, be an overestimate as a result of measurement problems arising when the Diagnostic Interview Schedule is used in follow-up studies.[26] An advantage of the data source for this project (the NPHS) over prior studies is that it provides an estimate of incidence and prevalence based on a comparable measure.

The predicted episode duration seems consistent with previous reports. Some community studies evaluating the duration of depressive episodes have reported comparable durations of episodes. For example, the mean duration of major depression in the Lewinsohn et al. adolescent sample was 23.6 weeks.[23] However, other studies have reported more brief average episode durations. For example, Rao et al.[27] reported an average episode duration of 10.3 weeks in a sample consisting of 17- and 18-year old women. Data from the Baltimore ECA follow-up study found that the median episode duration in that sample was only 12 weeks[28] and Kendler's follow-up of a community sample of female twin pairs reported a median time to recovery of only eight weeks.[29] These values for median duration are brief relative to the model-based estimate of mean episode duration from this study. This may relate to a right skew in the distribution of episode durations. Community studies have confirmed that a proportion of persons with major depression experience very protracted episodes. One analysis of ECA follow-up data found that 23.6% of subjects with major depression at baseline remained depressed after one year.[30] A study following a series of 78 clinical patients with major depression (probably with more severe and complicated disorders than those in the community studies) found that only 34 (48.6%) had recovered after one year. These findings suggest that most episodes of major depression in the community are brief (hence, having median durations less than three months), but that a proportion of persons experience very prolonged episodes, causing the mean duration to be greater than the median. In Kendler et al.'s study, mean episode duration was twice as long as median duration.[29] Any study relying on a finite and relatively brief follow-up interval will truncate the observed duration of more prolonged episodes.

Some of the differences between the duration of episodes reported by NPHS subjects and the mean durations estimated indirectly using incidence and prevalence data may simply reflect the concept of 12-month period prevalence, as measured in the NPHS. A person with the new onset of a protracted episode in the few weeks preceding the NPHS interview would have their duration of depression recorded only as the interval between onset and interview. This would also occur for a protracted episode that resolved a few weeks into the year preceding the NPHS interview. Finally, whereas some episodes of this condition can last for years, the maximum value that could be recorded using the questions employed in the NPHS was 52 weeks. Hence, the distribution of episode durations might have been substantially truncated.

Elements of consistency between these results and published findings have been summarized in the preceding paragraphs. It should be emphasized that the validity of a result is not confirmed by its consistency with other findings. Nevertheless, the development of dynamic descriptions of major depression epidemiology will need to be guided by principles of consistency with the existing literature of epidemiological studies.

The NPHS data appear to suggest that a substantial proportion of persons experiencing the major depressive syndrome will have a relatively brief disturbance. In itself, this has important public health implications. For example, a large number of false positives are likely to emerge from screening efforts if instruments like the CIDI short form are used to detect people in need of clinical intervention. However, these results serve to emphasize that the duration data from the NPHS interview do not necessarily reflect the average experience of persons with major depression. This appears to be a heterogenous condition, characterized both by brief and protracted episodes.

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Emigration Patterns of Cancer Cases in Alberta, Canada

Juanita Hatcher and Marilou Hervas

Abstract

Cancer registries are a unique source of data for population-based analysis of survival of cancer cases, but information on current vital status is essential. This paper describes a method to determine the last known vital status of cases and the emigration pattern of cancer cases diagnosed in Alberta. Data from the Alberta Cancer Registry (ACR) for the years 1985–1993 (83,446 cases) were linked to the Alberta Health Care Insurance Plan (AHCIP) registration file to identify cases that had left the province and the date they emigrated. Ninety-nine percent of the ACR cases linked correctly to the AHCIP registration file. Three percent of cases had left Alberta by March 1998. For the first five years of follow-up between 0.6% and 0.8% of cases alive at the beginning of each year of follow-up left the province in the succeeding year. Seven percent of those diagnosed under 45 years of age left the province compared to less than 2% of those aged 65 and over. There was no difference in emigration patterns between the sexes. The cancer sites with good prognosis tended to have the highest proportion of emigrants.

Key words: *cancer; emigration*

Introduction

The primary function of cancer registries is to identify and register all incident cancer cases occurring within their jurisdiction and to record information related to the death of each cancer case.[1] Cancer registries are thus a unique source of data for analyzing the survival of the population of cancer cases.[2] As many cancer registries do not actively follow up the incident cases, however, the residency and current vital status of those not known to be dead are often unknown, as is the impact of emigration on survival.

In Alberta, Canada, the problems caused by lack of active follow-up were overcome by linking the Alberta Cancer Registry (ACR) to the Alberta Health Care Insurance Plan (AHCIP) registration file maintained by Alberta Health and Wellness (AHW). This enabled the ACR to identify those cancer cases that had left the province and the date that they left. This date would be used as a censoring date in survival analysis, permitting the determination of the extent of emigration of cancer cases from Alberta and the potential impact on survival analysis.

This paper describes the pattern of emigration of cancer cases diagnosed in Alberta.

Data

AHW maintains a historical database of all residents of the province of Alberta who are/were registered in AHCIP, which insures all required medical care available in the provinces. The contract holder pays the premiums, and benefits are provided for all members of the immediate family. All residents of Alberta, except serving members of the Royal Canadian Mounted Police and the Canadian Military, inmates of federal penitentiaries and Status Indians, are eligible to register in AHCIP. The federal government pays the AHCIP premiums for Status Indians and other exceptions who are included in the AHCIP registration file. Only 200–300 of the approximately 3 million who are eligible to register choose not to. The AHCIP registration file includes data on more than 99% of the Alberta population.

The AHCIP registration data include a unique personal identifier, the Personal Health Number (PHN) and/or the AHCIP number, surname, initials, date of birth, gender and postal code, as well as the date and reason for any changes in coverage by AHCIP. Prior to 1994, the individual identifier (AHCIP number) consisted of an eight-digit base number to identify the contract, and a three-digit individual identifier. If a person changed the contract under which he or she was covered, by, for example, leaving the family home to

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marry, his or her AHCIP number would change. The PHN, introduced in 1994, is unique to an individual and remains the same for life. AHW has assigned a PHN to people who were initially registered before 1994, and a conversion file exists to link pre-1994 AHCIP numbers to the new PHNs. AHW is usually informed of people who leave the province by the medical plan of the province to which the person has emigrated. The AHCIP file is linked to the vital statistics file on a weekly basis to identify those who have died in the province. The estate of a dead person may also notify AHW when it receives the bill for the premiums.

The ACR is mandated under the Alberta *Cancer Program Act* to capture information on every incident of invasive cancer diagnosed in Alberta.[3] Pathology reports and death information from the Alberta Registries vital statistics file are the two main sources for identifying incident cases of cancer in Alberta. The ACR includes identifying information, (PHN, AHCIP number, current surname, previous surnames, first names, date of birth, gender and postal code); information on the incident tumour; and for those cases known to have died, date and cause of death. The Alberta Registries vital statistics file also provides information on cases that have died in Alberta. The ACR would not generally be informed of the deaths in other provinces. There has been no follow-up information on cases that are not known to have died.

Methods

All invasive cancer cases diagnosed in Alberta between January 1, 1985, and December 31, 1993, inclusive were identified from the ACR. A file containing PHN, AHCIP number, surname, initials, gender and date of birth for each individual diagnosed with cancer was submitted to AHW. This file included records for residents and non-residents of Alberta. A hierarchical deterministic linkage strategy was used to link the ACR to the AHCIP registration file, linking on the AHCIP number, the AHCIP base number and surname, gender and month and year of birth (identical). All links were checked using the other identifying information common to the two files.

Discrepancies between the ACR and the AHCIP records were investigated further by reviewing the patient's chart. Errors in the ACR were corrected, and AHW was notified of suspected errors in its data set. The agreement in the vital statistics of the linked cancer cases in the two files was also checked. Where the ACR had recorded that a case had died, but the AHCIP record indicated the case was alive, the death information held by the ACR was confirmed. In cases where AHCIP indicated the case had died but the ACR had no record of death, the case was deemed to be dead for subsequent analysis and the death date taken as that recorded by AHCIP. The ACR also had death information on some cases that AHCIP deemed had left the province.

Cases for which the first invasive primary (excluding non-melanoma skin cancer) occurred between 1985 and 1993 and which were resident in Alberta at the time of diagnosis were identified from the linked cohort. The percentage of cases that had left the province as determined from AHCIP was examined for each cancer site. Crude, age specific and age-standardized percentages were estimated. The standard population used was the total population of cancer cases used in the analysis. The percentage emigrating each year up to five years past diagnosis was estimated using the revised ACR vital status definitions. Person-years at risk were calculated as the time between diagnosis and date of last follow-up, death, emigration or for those still alive and resident in Alberta, as of March 31, 1998. All follow-up periods were terminated at five years to standardize the potential length of follow-up. The reduction in person-years at risk produced by censoring emigrants at the time of emigration rather than as of March 31, 1998, was investigated.

Results

Of the 83,446 cases that were resident in Alberta for at least one diagnosis between 1985 and 1993, 82,466 cases (98.8%) linked correctly to the AHCIP registration file. There was no difference in linkage rate between those who were registered as dead (98.8%) and those who were not registered as dead (98.8%).

The following results are presented for the 82,466 cases who were residents in Alberta at the time of diagnosis between 1985 to 1993, and who linked successfully with the AHCIP registration file (Table 1). Of the 45,925 cases that were registered as dead on the ACR, 97.3% were also registered as dead on the AHCIP registration file, 1.8% were registered as still alive and 0.9% were registered as having left the province. Of the 36,541 cases that were not registered as dead on the ACR, 1.3% were registered as dead on the AHCIP registration file and 5.6% had left the province. Of the 44,689 cases who were confirmed as dead on both the ACR and the AHCIP registration file, 574 (1.3%) of the dates of death do not agree between the two files. In 95% of these cases the AHCIP registration death date is later than that of the ACR. The discrepancies can be quite large: only 31% have differences of less than one year.

The following results are presented for the 76,164 cases of a first primary invasive cancer (excluding non-melanoma skin cancer) diagnosed between 1985 and

TABLE 1
Agreement between linked records for vital status on ACR and AHCIP files

ACR vital status	AH vital status			
	Dead	Alive	Left Alberta	Total
Dead	44,689 (97.3%)	836 (1.8%)	400 (0.9%)	45,925 (55.7%)
Alive	478 (1.3%)	34,012 (93.1%)	2,051 (5.6%)	36,541 (44.3%)
Total	**45,167 (54.8%)**	**34,848 (42.3%)**	**2,451 (3.0%)**	**82,466**

TABLE 2
Vital status of cancer cases diagnosed by cancer site in Alberta 1985–1993 as determined from 1998 AHCIP files

	Dead	(%)	[adj[a] %]	Alive in AB	(%)	[adj[a]%]	Left AB	(%)	[adj[a] %]	Total
Prostate	4,322	46.4	40.5	4,816	51.7	57.6	177	1.9	1.9	9,315
Female breast	3,383	31.3	34.1	7,007	64.8	62.4	418	3.9	3.5	10,808
Lung	8,634	87.9	85.7	1,029	10.5	12.4	161	1.6	1.9	9,824
Colorectal	4,999	57.0	53.9	3,531	40.3	42.9	235	2.7	3.2	8,765
Melanoma skin	428	19.1	29.7	1,668	74.5	66.2	142	6.3	4.1	2,238
NHL[b]	1,374	55.0	57.6	1,039	41.6	39.3	87	3.5	3.1	2,500
Leukemia	1,392	58.5	61.3	911	38.3	35.9	78	3.3	2.8	2,381
Uterus	528	23.7	26.3	1,627	73.1	70.4	72	3.2	3.3	2,227
Bladder	1,191	44.6	40.3	1,368	51.3	54.8	109	4.1	4.9	2,668
Kidney	940	48.1	50.4	929	47.5	45.4	86	4.4	4.1	1,955
Testis	34	5.1	40.8	571	85.9	56.0	60	9.0	3.2	665
Cervix uteri	394	31.6	48.2	771	61.9	48.0	81	6.5	3.7	1,246
Pancreas	1,968	95.3	92.9	79	3.8	6.2	19	0.9	1.0	2,066
Other	11,161	63.8	67.0	5,812	33.2	30.4	533	3.0	2.7	17,506
Total	**40,748**	**54.9**		**31,158**	**42.0**		**2,258**	**3.0**		**74,164**

[a] Age adjusted to the total cohort age distribution
[b] Non-Hodgkin's Lymphoma

1993 while resident in Alberta. On average, 3.0% of the cases had left Alberta by 1998. (Table 2) This is somewhat lower than the age-standardized five-year emigration rate from Alberta between 1991 and 1996 of 3.5%.[4] For the cohort of cases diagnosed between 1985 and 1993, and followed until March 1998, the cancer sites with the highest proportion of emigrants were testicular cancer (9.0%), cervical cancer (6.5%) and melanoma skin cancer (6.3%). The cancer sites with the lowest proportion of emigrants were prostate cancer (1.9%), lung cancer (1.6%) and pancreatic cancer (0.9%) (Table 2). Age standardization removes these differences except for those cancers with short survival rates, lung and pancreatic cancer, and for prostate cancer. In general, younger cancer cases were more likely to emigrate than older ones, but there is little difference in emigration rates among cancer sites. (Table 3).

For the first five years of follow-up, between 0.6% and 0.8% of all cases alive at the beginning of each year of follow-up left the province during the succeeding year. (Table 4). This pattern continues after five years of follow-up. There are no marked differences in the patterns of emigration among the cancer sites. There are no differences in emigration patterns between the sexes.

The person-years at risk for the first five years of follow-up is reduced on overall by 1.7% of the total person-years at risk if all persons not known to be dead are assumed to be alive. The largest effect (3.5% reduction) is seen for cancer of the testis, and the smallest (0.8% reduction) for cancer of the prostate.

Discussion

Survival analysis requires the follow-up of all incident cancer cases included in the analysis so that their current vital status is known. The ability of cancer registries to follow up cases is dependent on the health care system in which the registry operates, and the registration practice of each registry.[5] The main method of determining the death information for each cancer case is to link to the death certificate data for the jurisdiction of the registry. The ability to determine whether a case has left the jurisdiction of the registry and thus is lost to follow-up varies among registries. For the Scandinavian registries, complete follow-up is possible due to the existence of a unique lifetime identifier for each citizen. Thus date of emigration and date of death are known for all cancer cases.[5] In Saarland, Germany, follow-up is largely passive, due to restrictive legislation. It is possible to determine the death information for cases that die in Saarland, but not whether a case has emigrated.[5] In Canada, each of the provinces and territories operates its own registry and submits data to the national Canadian Cancer Registry. Most of these registries link their data to provincial vital statistics death information to determine the death information for those cases that die within the relevant province or territory. The Canadian Cancer Registry is linked to the National Death file for determining deaths among all cases that occurred anywhere in Canada. At the time of this study, the national death linkages have not been completed for the more recent years.

TABLE 3
Distribution by age and site of cancer cases who left Alberta after diagnosis 1985–1993

Age	0–44 yrs		45–64 yrs		65–74 yrs		75–84 yrs		85+ yrs		Total
Site	**Left/Diagnosed**	**%**	**Left/Diagnosed**	**%**	**Left/Diagnosed**	**%**	**Left/Diagnosed**	**%**	**Left/Diagnosed**	**%**	**Left/Diagnosed**
Prostate	0/10	0.0	60/1,924	3.1	69/3,892	1.8	38/2,772	1.4	10/717	1.4	177/9,315
Female breast	105/1,892	5.5	212/4,751	4.5	58/2,355	2.5	32/1,379	2.3	11/431	2.6	418/10,808
Lung	12/317	3.8	83/3,670	2.3	35/3,470	1.0	27/2,000	1.4	4/367	1.1	161/9,824
Colorectal	29/404	7.2	108/2,871	3.8	49/2,593	1.9	33/2,129	1.6	16/768	2.1	235/8,765
Melanoma skin	88/928	9.5	47/776	6.1	4/303	1.3	3/166	1.8	0/65	0.0	142/2,238
NHL[a]	40/493	8.0	25/910	2.7	13/578	2.2	8/413	1.9	1/106	0.9	87/2,500
Leukemia	40/640	6.3	16/654	2.4	12/524	2.3	9/399	2.3	1/164	.6	78/2,381
Uterus	8/155	5.2	39/1,005	3.9	13/675	1.9	11/332	3.3	1/60	1.7	72/2,227
Bladder	23/172	13.4	52/876	5.9	22/832	2.6	10/611	1.6	2/177	1.1	109/2,668
Kidney	22/265	8.3	45/795	5.7	13/520	2.5	6/294	2.0	0/81	0.0	86/1,955
Testis	56/584	9.6	4/73	5.5	0/7	0.0	0/1	0.0	0/0		60/665
Cervix uteri	58/659	8.8	20/350	5.7	2/143	1.4	0/67	0.0	1/27	3.7	81/1,246
Pancreas	1/66	1.5	3/614	0.5	7/646	1.1	7/511	1.4	1/229	0.4	19/2,066
Other	262/3,564	7.4	169/5,761	2.9	66/4,174	1.6	25/2,940	0.9	11/1,067	1.0	533/17,506
Total	**744/10,149**	**7.3**	**883/25,030**	**3.5**	**363/20,712**	**1.8**	**209/14,014**	**1.5**	**59/4,259**	**1.4**	**2,258/74,164**

[a] Non-Hodgkin's Lymphoma

TABLE 4
Distribution of time from diagnosis to leaving Alberta, by cancer site, for cancer cases diagnosed in Alberta 1985–1993

	0–<1 yr			1–<2 yrs			2–<3 yrs			3–<4 yrs			4–<5 yrs		
Interval	**# alive at beginning of interval**	**# leaving in interval**	**% leaving in interval**	**# alive at beginning of interval**	**# leaving in interval**	**% leaving in interval**	**# alive at beginning of interval**	**# leaving in interval**	**% leaving in interval**	**# alive at beginning of interval**	**# leaving in interval**	**% leaving in interval**	**# alive at beginning of interval**	**# leaving in interval**	**% leaving in interval**
Site															
Prostate	9,309	31	0.3	8,302	28	0.3	7,402	30	0.4	6,014	26	0.4	4,731	17	0.4
Female breast	10,801	51	0.5	10,161	73	0.7	9,415	66	0.7	8,084	52	0.6	6,702	37	0.6
Lung	9,821	71	0.7	3,367	28	0.8	1,929	22	1.1	1,315	9	0.7	960	13	1.4
Colorectal	8,763	45	0.5	6,477	51	0.8	5,385	42	0.8	4,292	28	0.7	3,412	13	0.4
Melanoma skin	2,235	19	0.9	2,138	29	1.4	2,022	31	1.5	1,741	16	0.9	1,441	13	0.9
NHL[a]	2,498	19	0.8	1,788	17	1.0	1,518	18	1.2	1,232	6	0.5	999	8	0.8
Leukemia	2,380	18	0.8	1,637	16	1.0	1,395	15	1.1	1,152	8	0.7	935	7	0.7
Uterus	2,227	12	0.5	2,050	18	0.9	1,932	9	0.5	1,704	5	0.3	1,480	7	0.5
Bladder	2,667	16	0.6	2,252	21	0.9	1,994	19	1.0	1,763	11	0.6	1,567	9	0.6
Kidney	1,954	17	0.9	1,401	13	0.9	1,247	15	1.2	1,047	10	1.0	862	13	1.5
Testis	665	9	1.4	640	14	2.2	615	5	0.8	568	7	1.2	501	3	0.6
Cervix uteri	1,244	14	1.1	1,089	11	1.0	951	16	1.7	821	10	1.2	707	9	1.3
Pancreas	2,066	17	0.8	294	2	0.7	123	0	0.0	77	0	0.0	52	0	0.0
Other	17,499	142	0.8	10,221	102	1.0	8,187	83	1.0	6,728	46	0.7	5,515	47	0.9
Total	**74,129**	**481**	**0.6**	**51,817**	**423**	**0.8**	**44,115**	**371**	**0.8**	**36,538**	**234**	**0.6**	**29,864**	**196**	**0.7**

[a] Non-Hodgkins Lymphoma

TABLE 5
Person-years of follow-up with and without censoring at time of emigration for cancer cases diagnosed in Alberta, 1985–1993

Site	Total person-years of follow-up		Difference	%
	Not censored	Censored		
Prostate	33,099	32,829	270	0.82
Female breast	43,165	42,542	623	1.46
Lung	11,852	11,673	179	1.54
Colorectal	25,332	24,988	344	1.38
Melanoma skin	9,339	9,078	262	2.88
NHL	7,248	7,079	169	2.38
Leukemia	6,630	6,523	108	1.65
Uterus	9,026	8,898	129	1.45
Bladder	9,726	9,550	176	1.84
Kidney	5,945	5,801	144	2.48
Testis	2,982	2,883	100	3.47
Cervix uteri	4,639	4,497	142	3.15
Pancreas	1,218	1,198	20	1.65
Other	41,137	40,218	919	2.28
Total	**211,338**	**207,755**	**3,583**	**1.72**

To overcome this problem, the ACR sought other solutions to determine the current residence and vital status of those cancer cases not known to have died. The AHCIP registration file, which AHW maintains, provided ACR with the required information. Deterministic linkages were used in preference to probabilistic linkages because of the availability of the AHCIP number. Evaluation of the linkage confirmed that it provided a high degree of accuracy (98.8% among cases resident in Alberta). However, this information is subject to the problems inherent in using administrative data in an application for which such data were not originally planned.[6] Some of these problems have been identified, but not resolved, in this study. There are small percentages of cases in which the vital status is different in the two files (1.6%), in which the recorded residency patterns do not seem to concur with the stated residency in the ACR (1.6%), and in which the death dates do not agree (1.7%). The discrepancy in vital status could be explained by the potential delay in registering deaths on the AHCIP registration file for those cases that are dead in the ACR data and alive in the AHCIP data. For those cases that are alive in the ACR data and dead in the AHCIP data, the absence of the PHN on the ACR record may have prevented appropriate linkage to the Alberta Registries vital statistics death data, or these may be Alberta residents who died outside the province. These cases would not be included in the Alberta Registries vital statistics computer files. However, AHW may have been notified of the death when payment to AHCIP ceased.

The main aim of the project was to identify the residency and vital status of those cases that were not identified as dead so that cases leaving the province could be censored at the date of leaving in any survival analysis. The linkage has shown that cancer cases do in fact move out of the province, although they are less likely to migrate than the general population.[4] The probability of their emigration depends on the site of their cancer, their age, and the time since diagnosis. As would be expected, those who are diagnosed at a younger age, and are diagnosed with a site for which there is a good prognosis, are more likely to move out of the province. The low emigration rate (1.9 %) for prostate cancer, which has a relatively good prognosis, may in part be explained by the advanced age at diagnosis. Both the good prognosis and the younger age at diagnosis can explain the high emigration rates for cancer of the testis, cervix and melanoma. Among those cases that migrate, the proportion of live cases that migrate on an annual basis does not vary appreciably among the cancer sites.

The implications of these findings on the results of survival analysis may be sizeable depending on the reasons that the cases migrate. If the cases that migrate were the ones for whom the prognosis is good compared to others with the same diagnosis, the survival rates would tend to be underestimated. However, if the ones who leave have poor prognoses compared to others with the same diagnosis then the survival rates would be overestimated. The cause-specific survival of foreign residents of Geneva diagnosed with cancer is superior than that of native residents, which may be due to a combination of the healthy immigrant bias and/or the repatriation of those cases with poor prognosis (unhealthy emigrant bias).[7] The Alberta data show that the proportion of cases migrating from Alberta tends to be related to the overall prognosis of the diagnosis. The ACR has not collected staging information, and therefore is not able to address the issue of the prognosis of those cases who emigrate relative to those who remain in Alberta. However, the similarity of the distribution of time to emigration among the cancer sites for those who migrate would not indicate any systematic emigration patterns within site, based on prognosis.

Censoring of the emigrants at the time they leave the province decreases the overall person-years at risk by 1.7%. Although this figure is small, it may have a marked effect on survival estimates, particularly where survival is short or emigration is high. In the Eurocare II study, active follow up of lung cancer cases not known to be dead after five years resulted in decreases in five-year survival of up to 2.5%.[5]

Alberta is able to identify the cases who emigrate because the AHCIP registration file is updated regularly, in part to ensure that appropriate health care premiums are paid. In other provinces, the health care registration

file may either not be available to the cancer registry or may not be sufficiently up to date. Thus the national death clearance which is currently being undertaken is essential for provincial cancer registries to improve their survival analysis. However, this death clearance will not identify cases who die outside of Canada and thus linkage with the AHCIP registration may still be of value.

The economy in Alberta is largely dependent on the oil and natural gas industry, which fluctuates with world economic conditions. The economy tends to drive the emigration pattern for the younger age group, while that of the older age groups may be driven by the desire to seek more clement winter conditions in other provinces or jurisdictions. Given these patterns of emigration, the results may not be applicable to other jurisdictions, but do indicate that emigration of cancer cases is an issue that should be addressed when undertaking survival analysis.

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The Economic Burden of Mental Health Problems in Canada

Thomas Stephens and Natacha Joubert

Abstract

This study provides a comprehensive estimate of the economic burden of mental health problems in Canada in 1998. In particular, it estimates the cost of non-medical services that have not been previously published and the value of short-term disability associated with mental health problems that was previously underestimated, according to the approach used here. The costs of consultations with psychologists and social workers not covered by public health insurance was $278 million, while the value of reduced productivity associated with depression and distress over the short term was $6 billion. Several data limitations suggest that these are underestimates. The estimated total burden of $14.4 billion places mental health problems among the costliest conditions in Canada.

***Key words**: Canada; depression; distress; economic cost; population*

Introduction

The objective of this study is to provide a comprehensive estimate of the economic burden of mental health problems in Canada. In so doing, we seek to build on estimates published in Health Canada's *Economic Burden of Illness in Canada*, *1993* (EBIC, 1993)[1] and to address some of the data issues identified in this complex analytical task. While direct and indirect economic costs are only one aspect of the burden of disease, they can provide a valuable perspective for planning programs and setting priorities.

A recent study by Health Canada's Cancer Bureau of the former Laboratory Centre for Disease Control (LCDC), estimates that the economic burden of mental disorders in Canada was $7.8 billion in 1993,[1] or $8.4 billion in 1998 dollars. Mental disorders ranked seventh among the 20 disease categories for which cost estimates were published. *Direct* costs for treating medically diagnosed mental disorders totalled $6.3 billion (1998), comprising $3.9 billion for hospital care, $887 million for other institutional care, $854 million for physician care, and $642 million for prescription medications. Additional *indirect* costs totalling $3.0 billion were made up of short-term sick days ($866 million), long-term disability ($1,707 million), and premature death ($400 million), although these latter amounts were not restricted to diagnosed disorders.

These estimates were based on a societal perspective and thus incorporated both direct (internal) and indirect (external) costs, using conventional assumptions for the calculations. For example, the value of lost productivity due to early retirement was based on the present value of the lifetime earnings of the person who retires early due to a mental disorder. While there are some limitations to the approach (for example, health care *savings* arising from early death were not considered), it is consistent across disease categories and allows for a reasonably fair comparison of the economic burden of diseases.

However, with respect to the economic burden of mental health problems in particular, there are some more serious limitations to this approach. First, it includes only *medically treated*, *diagnosed* disorders in the direct costs (ICD-9 codes 290–319). By definition, these are problems that come to the attention of the health care system and that do not include states such as distress or depression that are untreated by physicians or other health professionals providing publicly insured health services.

Large numbers of Canadians with mental health problems treated outside the medical system are missing in such medically based calculations of the direct cost of illness. According to data from the 1996/97 National Population Health Survey (NPHS),[2] only 21% of Canadians who consulted a psychologist about their mental health also consulted a family doctor or a psychiatrist in the previous year, while 29% who consulted a social worker also consulted a physician. Since 4% of Canadians were depressed and 20% were

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classified as distressed in 1996/97,[3] the direct costs associated with their mental health problems could be considerable, but most of these would not be included in the EBIC, which is based on publicly insured services.

A second limitation of the EBIC, which affects the *indirect* cost of mental health problems, is the method used to attribute short-term disability to specific disease categories. Unlike the direct costs, the indirect costs are not limited to diagnosed disorders in the EBIC analysis, but include any health-based reason for cutting down on normal activity. The attribution to specific disease categories is then made on the basis of data from the Quebec Health Survey.[1] Although the Quebec data are the only available basis for attributing short-term disability to disease categories, they have important shortcomings with respect to their validity and applicability. First, the validity of reports attributing activity reduction to mental health problems is questionable because a significant proportion of these attributions is based on third-party reports for other members of the household. Second, even if these reports were of unquestioned accuracy, the application of these 1992/93 Quebec data to all of the Canadian population is doubtful: in 1994/95, Quebec residents were the *least* likely to report that distress affects their life – 13% versus an average of 17% for the other provinces.[3] This low level of attribution of effects to distress in Quebec leads to underestimating the short-term disability costs of mental health problems such as distress, estimated in the EBIC, 1993 as $811 million ($866 million in 1998 dollars).

The objective of this analysis is to address these shortcomings and provide a more complete estimate of the economic burden of mental health problems in Canada. In so doing, we are building on one of the recommendations in the EBIC, 1993,[1] namely to "improve data sources and refine methods for direct and indirect cost components to provide more comprehensive information for specific diseases" (p. iv).

Methods

Data source

The source of data for the original analyses in this study was the 1996/97 NPHS "share" file. The share file is virtually identical to the public use file, but includes some detail removed from the latter for reasons of respondent confidentiality. Population estimates and dollar values were adjusted to 1998 figures.[4,5]

The NPHS is the biennial survey conducted by Statistics Canada to describe health status and health determinants; the 1996/97 sample is representative of the household population of Canada. Data collection for the mental health indicators in the present study was by personal interview of approximately 77,000 persons aged twelve and over.[2]

Definitions of mental health problems

We used the NPHS questions on depression and distress as evidence of mental health problems. The distress scale includes many symptoms of anxiety (e.g., feeling nervous, restless or fidgety) and, with the depression scale, provides a reasonably comprehensive view of population mental health problems. Depression was defined according to the Statistics Canada definition[2] as a probability of 90% or greater of a major depressive episode in the previous year; the overall prevalence rate is 4%. Unlike depression, there is no independently verified definition of "high distress" for the measure used in the NPHS. We used as a definition a response of "a lot" or "some" to the question "How much do these (distressing) experiences interfere with your life or activities?" regardless of level of distress on the 24-item scale preceding the question on impact. By this definition, 15% of Canadians can be regarded as distressed.

There is a fairly high association between depression and distress: 53% of depressed persons also reported distress, and 24% of distressed persons were depressed. In order to avoid double-counting these persons, all analyses in this paper consider two groups in turn – all depressed persons, then distressed persons who are free of depression.

Direct costs

The EBIC, 1993 uses a "top-down" approach to estimating the direct costs of illness. That is, estimates are based on a known total for health care costs, which is then allocated to various disease categories, according to the principal diagnosis for the care received. In contrast, this study is obliged to use a "bottom-up" approach, estimating the volume of non-medical health care associated with mental health problems, and then the associated cost.

The NPHS ascertained the number of consultations with each of psychologists, social workers, physicians, and other health professionals in the previous 12 months, for reasons of "physical, emotional or mental health." The survey also separately identified visits to psychologists, social workers, physicians and others, for reasons of one's "emotional or mental health," but did not ascertain the *number* of these mental health visits. To estimate the *number of social worker and psychologist visits for mental health reasons,* we combined data from these two separate questions. Further, to exclude publicly insured consultations with psychologists and social workers (e.g., in hospitals) already included in EBIC estimates, in the absence of survey data on the *location* of the consultation, we adjusted the total of visits to reflect the proportion provided by psychologists or social workers *in the absence of any physician consultation*. As noted above, this is 79% of those consulting a psychologist and 71% of persons seeing a social worker.[a]

[a] For psychologists, this corresponds reasonably well to an estimate of 69% of service hours spent in private, as distinct from institutional, settings by 1,065 respondents to a 1999 survey of the 3,240 psychologists registered with the Canadian Register of Health Service Providers in Psychology.[6]

The number of psychologist and social worker visits by all depressed and distressed/not depressed persons was obtained from the NPHS, adjusted for the proportion of institutional visits and further adjusted for a population growth of 1.4% between 1996/97 and mid-1998,[4] and then multiplied by the average cost of such visits ($125).[b]

With respect to medications, there are severe limitations to the NPHS data. They are restricted to reports of any use in the previous month and there is no information that would permit an estimate of the *annual frequency* of use. Unless frequency can be obtained from some other recent and comparable source, the cost of these medications has to be limited to the cost estimated in EBIC, 1993, which is confined to prescriptions arising from medical care.

Indirect costs

The indirect costs of mental health problems not fully accounted for in the EBIC, 1993 analysis are, for reasons described above, those due to short-term work loss associated with depression or distress. In the present study, short-term work loss is calculated from the NPHS using the questions on two-week disability days (cut-down days + bed-days). *Excess* time off associated with depression is obtained by comparing the disability days of depressed vs. non-depressed persons and then, similarly, distressed vs. non-distressed persons. Although the exact health reason for the time off was not ascertained and has to be assumed to be mental-health related, this is analogous to the procedure that attributes excess sick-days to smokers.[7]

Since the prevalence of depression and distress varies according to labour force status, we estimated work-loss days separately for part-time workers, full-time workers, and non-employed persons. As the NPHS does not identify on which days of the week the reduced activity occurred, the proportion that are work days was estimated by assuming that the probability is equal for any day of the week being a sick day, and multiplying the total by 5/7 (usual work days/week). Assuming two weeks of annual holidays, the two-week total was then multiplied by 25 to give an annual estimate for work-days of restricted activity. In the absence of an exact report of hours worked per week, we weighted the work-loss of part-time workers by a factor of 0.5 in estimating their contribution to the total for the worker population. The dollar value of this lost time was calculated using average employment income for full- and part-time workers as published by Statistics Canada,[8] expressed in 1998 dollars.

To maintain consistency with the approach used in EBIC, 1993, the disability-days of persons outside the work force was also calculated. A proportion of the disability days of full-time workers (two of seven days) and of part-time workers (four and a half of seven days) was added to this total, to account for their activity restriction outside of the usual working days. The value of this lost time was obtained by assuming unpaid work is worth $15,000 annually, based on an hourly wage of $7.50 and 2000 hours of work annually. This is consistent with the "generalist" approach to the value of unpaid work as used in EBIC, 1993.[9]

Next, total disability days were adjusted for the fact that most are not days of complete inactivity, but only of reduced activity. For working persons, such "cut-down" days constitute 74.1% of all two-week disability days.[2] If cut-down days are weighted as 0.5 of a bed-day, then the adjustment required to take account of the proportion of cut-down days is $(74.1 \times 0.5 + [1 - 74.1] \times 1.0) = 0.6285$.

Results

Direct costs – visits to non-medical mental health professionals

In 1996/97, depressed persons age 12 and older who sought professional help for mental health reasons made almost 1.5 million visits to social workers and more than 850,000 visits to psychologists (Table 1). The 1998 equivalent is estimated at 2.38 million visits in total, after adjusting for population growth of 1.4%.[4] In addition, 1.6 million Canadians reported being distressed without being depressed. While the vast majority of them did not seek care from any mental health professional, there were approximately 280,000 visits to social workers and 328,000 visits to psychologists. These consultations are the equivalent of 616,000 visits in 1998.

For depression and distress combined, there were almost 3 million visits to psychologists and social workers in 1998. An estimated 2.2 million of these visits took place on a fee-for-service basis outside institutions (Table 1). At $125 each, the total cost for these visits exceeds $278 million.

Indirect costs – days off work

In 1998, almost 678,000 employed Canadians accumulated more than 39,000 excess person-years of short-term reduced activity associated with depression and another 2 million had over 115,000 person-years of time off associated with distress (Table 2). Among unemployed Canadians, there were more than 76,000 days of reduced activity associated with depression and 224,000 associated with distress.

After adjusting for part-time work, inflation, and the preponderance of cut-down days over bed-days, the total value of lost work time was $2.16 billion. An amount equivalent to $3.86 billion in unpaid work was similarly reported by depressed and distressed persons. The total value of paid and unpaid work lost associated with these conditions was $6.02 billion in 1998 (Table 2).

[b] Provincial bodies that license psychologists and social workers were contacted for their fee schedules: six replied, representing 85% of the Canadian population. While fees range widely ($60-$180/session) among and within provinces, the weighted average is $125. This figure was confirmed as a reasonable estimate by the Canadian Register of Health Service Providers in Psychology (P. L-J. Ritchie, personal communication, October 13, 2000).

TABLE 1
Number of consultations with social workers and psychologists for reasons of mental health, Canada, age 12+, 1998

Condition	Social worker[a]	Psychologist[a]	Total[a]	1998
Depressed	1,491,423	858,223	2,349,646	2,382,541
Distressed (not depressed)	279,634	327,604	607,238	615,739
Both conditions				
– all settings	1,771,057	1,185,827	2,956,884	2,998,280
– fee basis only	1,257,450	936,803	2,194,253	2,224,973

[a] Source of unadjusted data: National Population Health Survey, 1996/97 share file

TABLE 2
Indirect cost of depression and distress, Canada, age 15+, 1998

Condition	Population affected[a]	Excess time lost		Cost $
		Average days in 2 weeks[a]	Total person-years	
Depressed				
– paid work	677,625	0.84	39,075	451,676,778
– unpaid work[b]	536,221	2.00	76,393	967,268,150
Distressed/not depressed				
– paid work	2,043,168	0.82	115,397	1,711,976,531
– unpaid work[b]	2,341,064	1.34	224,126	2,892,781,577
Both conditions				
– paid work	2,720,793	0.83	154,472	2,163,653,309
– unpaid work[b]	2,877,285	1.51	300,519	3,860,049,727
Total	**5,598,078**	**1.28**	**454,991**	**6,023,703,036**

[a] Source: National Population Health Survey, 1996/97 share file
[b] Includes the value of unpaid work of full- and part-time workers while not at work

Summary

Table 3 summarizes the direct and indirect costs for depression and distress as estimated in the present study, and the direct costs and indirect costs for medically treated mental disorders as estimated in EBIC, 1993. Our estimate for the economic burden of mental health problems increases previously published estimates[1] by 71% – after adjustment for inflation between 1993 and 1998. The total in 1998 was $14.4 billion.

Discussion

This attempt to provide a comprehensive estimate of the economic burden of mental health problems reveals that previous estimates[1] may be far too low, primarily due to attributing too small a proportion of lost productivity to mental health problems. However, the current study may also be an underestimate of the true value, due to several limitations. The principal ones are:

- The mental health problems newly accounted for in this analysis are limited to depression and distress. These conditions are important, but they are not exhaustive; others such as phobias could not be accounted for, although some anxiety symptoms are part of the distress scale.
- The NPHS definition of depression is conservative: it counts only those persons who report feeling "sad, blue or depressed for two weeks or more in a row" during the past 12 months *and* whose responses to a symptom checklist indicate a probability of a major depressive episode during the past year of 90% or more. This definition would exclude anyone with transient feelings of depression; such persons might well have taken time away from work or other usual activities, however.
- It was not possible to estimate the cost of over-the-counter medications possibly used in response to depression and distress. As collected by the NPHS, these would be sleeping pills, painkillers, stomach remedies and laxatives, but person-level data on frequency of use needed to calculate annual consumption are not available.

TABLE 3
Summary of costs related to mental health problems, Canada, 1998 ($ million)

	Cost	Source[a]
Treatment		
– of diagnosed disorders		
– medications	642	EBIC
– physicians	854	EBIC
– hospitals	3,874	EBIC
– other institutions	887	EBIC
– of depression and distress		
– non-publicly insured mental health professionals	278	This study
Total	6,257	
Lost productivity		
– short-term disability	6,024	This study
– long-term disability	1,708	EBIC
– early death	400	EBIC
Total	8,132	
Total	**14,389**	

[a] EBIC estimates from Reference 1, adjusted for inflation of 6.68% between 1993 and 1998.[5]

- It is not clear that the true extent of reduced productivity due to distress and depression is captured by the question used in the NPHS, "[During the last 14 days], did you stay in bed/cut down on normal activities because of illness or injury?" It seems unlikely that distress, in particular, would be universally regarded as an "illness."
- Long-term work loss was not included in this analysis due to data limitations, including the strong possibility that depression and distress are the *result* as much as the *cause* of activity restriction.
- The value of reduced productivity of non-employed persons has been set at the equivalent of $15,000 per year – a very conservative figure.
- No account has been taken of the cost of violence and early school departure that may accompany depression and distress. Nor have we included the costs of smoking, drug and alcohol abuse that may be used to cope with depression and distress, nor the cost to family and friends of providing support to persons in need. Further, no estimate has been made here of the large amounts of time devoted to personal crisis counselling by other professionals, e.g., guidance counsellors in schools, EAP staff in work settings, and clergy in the community.

All of the foregoing limitations would produce underestimates of the true economic burden. Only one limitation – co-morbidity – might inflate these estimates.

While it is possible to estimate the excess days off work associated with depression and distress (Table 2), there is no way of knowing whether mental health problems are the *primary cause* of the lost productivity, or whether a co-existing condition might be the cause, since the NPHS does not provide this detail (and, as noted, other sources such as the Quebec Health Survey are inadequate for this purpose). If conditions other than mental health problems lie behind the reduced productivity associated with depression and distress in this analysis, however, they appear to be limited: depressed persons report *half* the number of co-existing physical conditions (1.8) of persons with chronic physical conditions (3.3) (G. Torrance, personal communication, May 1, 2000). While there is an association between depression and number of physical health problems, and similarly between distress and physical health, it is modest for those with one or two physical conditions.[3]

The upshot of all these limitations is that the estimates presented in this paper are likely quite conservative. We can thus conclude with fair confidence that the economic burden of mental health problems – both medically treated and not – is $14.4 billion annually, *at a minimum.*

Implications

The major implications of this study are similar in many ways to those described in a recent analysis of the mental health status of the Canadian population.[3] The new element is the dollar figures.

These results strongly suggest that promoting the mental health of Canadians would be a sound investment, not only to prevent mental health problems but also to reduce the staggering economic burden associated with them. This analysis demonstrates that these are much higher than suggested by previous studies,[1] and indeed, are likely considerably higher than the available data suggests, due to the many limitations described above.

Moreover, the number of persons in distress may *increase,* in tandem with current trends in child poverty, income disparities, involuntary part-time work, single-parenting, youth unemployment, and declining expenditures on health, welfare and education.[3,10] It is striking that youth now exhibit the highest distress levels in the population, when they had the lowest levels 20 years ago.[3,10] This trend raises the possibility of lifelong problems for the current youth cohort, exacerbated by the sharp decline in the support provided by the community and the mental health system. This situation will continue to deteriorate as long as individual support networks and the broader social safety net are not repaired and maintained. Whatever benefits children and youth may get from various programs can be very difficult to sustain when there is a lack of support from the family and the community.[11] Social support can be increased by fostering the development of meaningful relationships in families and social environments – in schools, workplaces, the community and institutions.[12]

It is clear that offering only more "services" will not respond effectively to the population's mental health needs. Since approximately 60% of people with mental health problems do not receive care from a health professional, the apparent gap in services is simply too big to fill. What is evidently needed is a different kind of investment to promote the population's mental health. Generally speaking, this could take the form of developing individual and community resourcefulness, and promoting resilience among individuals of all ages.[13,14] The significant contribution of mental health problems to the global burden of disease is being addressed by a growing number of countries, including Canada, the United States, Australia, New Zealand, and the Member States of the European Union, all of which are developing national plans of action or other initiatives to promote the mental health of their populations.[11] The dollars thus invested would represent a small figure compared with the economic burden if nothing is done, judging by our analysis.

This analysis also has implications for research, especially for the collection of data in future population surveys. In order to calculate both direct and indirect costs adequately, it is essential to determine the respondent's assessment of (a) the reason for cutting down on normal activities, in a manner that can deal with co-morbidity, (b) whether any mental health services received were covered by public health insurance, and (c) the frequency and dosage of over-the-counter medications taken for mental health reasons. To enhance the validity of these reports, they should be obtained directly from the respondent; third-party reports should not be accepted. Even with these improvements in data, the economic burden of mental health problems will likely continue to be underestimated until they are reported as openly as are physical health problems.

Acknowledgments

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The Storage of Household Long Guns: The Situation in Quebec

Michel Lavoie, Lise Cardinal, Antoine Chapdelaine and Danielle St-Laurent

Abstract

This survey on the storage of household firearms in Quebec was conducted in 1994. At that time, 35% (175/504) of survey participants who kept long guns in their homes had failed to comply with Canadian firearm storage regulations. In most cases (85%; n = 149), this was because at least one stored long gun was found to be both operable and accessible. Thirty-seven per cent of participants stated that no one, including themselves, had used their firearm(s) in the 12 months preceding the survey. These findings point to two possible ways of dealing with long guns kept in the home: render these weapons inoperable or inaccessible, which would increase the level of compliance with the regulations, and dispose of those no longer in use. The results of this survey have never been published before, and constitute the only information of this kind with respect to Quebec.

Key words: *firearms; home; storage; survey*

Introduction

Between 1989 and 1997, an average of 1,252 firearm-related deaths were reported each year in Canada. Of these, 80% were suicides, 15% were homicides, 4% were "accidents" and, in 1% of cases, the cause was unknown.[1] Approximately one third of these deaths (30%) occurred in Quebec.[2] Most firearm-related deaths (at least three in four) in Quebec were linked to the discharge of long guns (shotguns and rifles) or, more rarely, handguns (pistols or revolvers).[3] A Quebec study of 425 cases of firearm-related suicide occurring between September 1 and September 31, 1996 indicated that 30% of the victims were not the owners of the gun and that in most cases the gun had not been safely stored.[4]

In 1992, Canada had the sixth highest rate of firearm-related deaths (rate per 100,000 inhabitants, adjusted for age) among 26 countries deemed to have a high gross national product (World Bank classification). That year, Canada's rate of firearm-related deaths was 4.31, compared to 14.24 in the United States, which had the highest rate, and 0.05 in Japan, which had the lowest.[5] In 1993, the direct and indirect costs associated with firearm-related deaths and injuries in Canada was estimated at $6.6 billion.[6]

Case-control studies conducted in the United States have shown that the presence of a firearm in the home increased the risk of firearm-related death for household members and relations: the members of households where firearms were kept had a 4.7-time greater risk of committing suicide[7] and a 2.7-time greater risk of being the victim of a homicide[8] than those living in households where no firearms were kept. The risk of suicide increased ninefold (9.0) in cases where the firearm was stored loaded (compared to homes without firearms), or threefold (3.0) when the weapon was kept under lock and key or rendered inoperable.[7] Members of households in which there were firearms were also 22 times more likely to die from firearm-related injuries (suicide, homicide or accident) than they were to kill an intruder in self-defence using a firearm.[9]

Many experts believe that an important factor in reducing the number of firearm-related deaths and injuries is to reduce the accessibility of firearms by reducing the number of firearms in the home or by storing household firearms more safely.[10,11] In Canada, the *Firearms Act* contains provisions pertaining to both these strategies.[12] Provisions determining the rules governing firearm storage have been in place since January 1993.[13] These provisions stipulate that firearms

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must be stored unloaded, rendered inoperable or inaccessible and, preferably stored separately from the ammunition.[13]

This paper examines the results of a 1994 survey on household firearm storage practices in Quebec[14] in the context of Canadian firearm storage regulations.[13] These results provide an estimate of the level of compliance with firearms regulations shortly after these regulations came into force. They have never been published before and constitute the only available information on the situation in Quebec.

Methodology

The target population was composed of Quebec residents who were 18 years of age or older and who owned at least one firearm that they stored at home as of September 1, 1994. The total survey population comprised 515 firearm owners drawn from a random sample of 4,654 households selected from Quebec telephone directories. These households were selected to reflect the demographics of the various administrative regions in Quebec. Only 17% (n = 792) of the households were found to include an adult who owned a firearm.

Each gun owner in these households was invited to participate in the survey. A total of 524 firearm owners agreed to take part, although nine of them ultimately chose to terminate their participation before the survey was completed. This represents a participation rate of 65% (515/792). Ninety-eight percent (n = 504) of participants indicated that they kept at least one long gun (shotgun or rifle) in their home, but only 7% (n = 36) indicated that they owned a handgun (pistol or revolver). Only the results pertaining to the 504 owners of long guns are presented here. The maximum margin of error for a sample of this size (n = 504) is plus or minus 4.4 %, based on an alpha threshold of 0.05.

The data were collected between September 1 and 13, 1994, or approximately one and a half years after the firearm storage regulations[13] of the Canadian *Firearms Act* came into force[12] (January 1, 1993). Data collection was entrusted to a well-known professional polling firm, Le Groupe Léger & Léger Inc. The data were obtained directly from long gun owners by means of telephone interviews conducted by professional, bilingual interviewers using a pre-tested, standardized questionnaire that appeared on a computer screen. The data were recorded as the interview proceeded, with each interview lasting approximately 12 minutes.

The goal was to describe the state of long gun storage in the context of Canadian firearm storage regulations.[13] These regulations stipulate that firearm storage must meet the three following criteria to be deemed safe: the firearm must be stored unloaded (first criterion); it must be rendered inoperable or inaccessible (second criterion); and the ammunition must be securely stored (third criterion). A long gun is considered to be unloaded (first criterion) if there is no ammunition in the cartridge magazine. It is rendered inoperable (second criterion) by means of a secure locking device or by the removal of a part which is essential to its operation, such as the bolt. In order to be inaccessible (second criterion), a firearm must be stored in a securely locked place that cannot readily be broken open or into. This can be a room, a receptacle or a container. The ammunition (third criterion) must be stored separately from the firearm, or together with the firearm, provided that the place in which it is stored is inaccessible (securely locked and not readily broken into). In the latter case, the place can be a container or a receptacle but not a room.

Long gun storage practices were described by asking each participant about a single weapon. This was done for practical and methodological reasons, as several questions must be asked to evaluate the storage conditions for a single weapon. In cases where a participant indicated that he/she kept more than one firearm at home, one firearm was selected at random using a software program that made the selection from the list of firearms declared by the participant.

Participants were asked to respond to a series of questions specifically designed to determine whether the firearm was equipped with a secure locking device; whether any of its parts had been removed; whether it was being stored in a container, a receptacle or a room; whether the place of storage was locked and whether it could readily be broken open or into; whether the ammunition was stored with the weapon or in a separate place; and, finally, whether the weapon was stored loaded.

Each participant's answers were analyzed during the interview using a special software program to determine whether storage practices complied with the three criteria stipulated in the regulations. Then, each participant was placed into one of two categories: compliers (firearm storage complies with all three criteria) or non-compliers (firearm storage does not comply with at least one of the three criteria). In other words, the first group supposedly comprised those who stored their firearms securely, while the second group comprised those who failed to meet the safe storage criteria.

Findings

The majority of the participants were between the ages of 35 and 54 (53%). There were nine times as many men (n = 465) as women (n = 50). In 96% of cases, the language spoken was French, and nine times out of ten the participant was not the sole occupant of the household. A greater percentage of participants were from rural areas (60%) than from urban centres (40%). The majority of participants (64%) indicated that they had completed 12 years or less of schooling.

On average, participants kept 2.7 long guns in their homes. Thirty-two per cent owned only one gun, 29% owned two, 18% owned three, and 21% owned four or more. The three most popular long guns were, in decreasing order, 12-calibre shotguns (55%), .22-calibre rifles (42%), and .410-calibre shotguns (24%). In 22% of cases, participants indicated that other persons had

TABLE 1
Reasons given by participants for owning a firearm[a]

Reason stated	%[b]
Hunting	87
Target practice	11
Gun collecting	7
Souvenir	5
Self-protection	3
Employment	2
Predators and other pests	1
Other	1

[a] Shotgun or rifle.
[b] The total exceeds 100% because some participants (n = 504) mentioned more than one reason for owning a firearm (n = 593).

access to their firearms. Thirty-seven per cent of respondents indicated that no one, not even themselves, had used their firearms in the 12-month period preceding the survey.

Hunting was the most frequently mentioned reason (87%) for owning a long gun (Table 1). A small percentage of participants owned a firearm for self-protection (3%) or to hunt predators or other pests (1%). At the time of the survey, 91% of participants stated that they had received training in the handling of firearms and 53% were aware of the existence of an act governing the storage of firearms in Canada.

The study findings show that, at the time of the survey, 65% of participants (n = 329) had at least one long gun that was securely stored in their home (in accordance with the three criteria). Specifically, almost all participants indicated that the weapon for which they were providing information was stored unloaded (99.6%; first criterion) and that the ammunition for this weapon was also securely stored (91%; third criterion). Seventy per cent of the gun owners surveyed stated that their weapon had been rendered inoperable or inaccessible (second criterion).

On the other hand, the results also show that 35% of participants (n = 175) had failed to comply with at least one of the three criteria of safe storage. The likelihood of non-compliance was greatest (Table 2) among participants whose spoken language was French ($p < 0.05$); those who owned at least one weapon that was accessible to others ($p < 0.05$); and those who were not aware that there is a law governing firearm storage in Canada ($p < 0.05$). However, the other variables considered were not associated with firearm storage practices ($p \geq 0.05$). These included age (18–34; 35–54; 55+); gender (male, female); place of residence (rural, urban); years of schooling (≤ 12, > 13); family income (< $40,000, ≥ $40,000); living alone (yes, no); the presence of children under the age of 18 (yes, no); the number of firearms owned (1, 2, 3, 4+); the reasons for owning

TABLE 2
Factors associated with compliance with firearm storage regulations[a]

	Non-compliers[b] (n = 175)	Compliers (n = 329)	Total (n = 504)	
Factors	**%**	**%**	**%**	**(n)**
Language spoken				
French	36[c]	64	100	(482)
Other	14	86	100	(22)
Aware of the existence of an Act				
Yes	27	73	100	(267)
No	43[c]	57	100	(160)
DK/NR	43	57		(77)[d]
Firearm accessible to one other person				
Yes	46[c]	54	100	(109)
No	32	68	100	(394)

[a] Shotgun or rifle
[b] Failure to comply with at least 1 of the 3 storage criteria.
[c] $p < 0.05$
[d] Not taken into account.

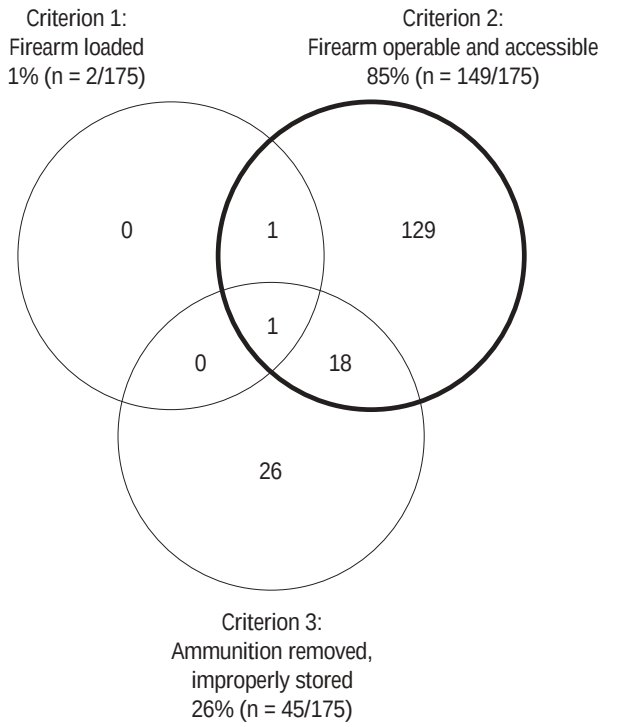
firearms (hunting, target practice, gun collecting); the calibre of weapon (12, .22, .410); weapon(s) not used during the past 12 months (yes, no); and prior training in the handling of firearms (yes, no).

Figure 1 represents the survey findings with respect to the 175 non-compliers. Eighty-five percent (n = 149) did not comply with Canadian firearm storage regulations by having at least one stored long gun which was both operable (not equipped with a secure locking device and comprising all of its parts) and accessible (stored in an unlocked place or a locked place that could readily be broken open or into).

The 149 non-compliers who kept at least one long gun that was operable and accessible (non-compliance with the second criterion of firearm storage) were questioned more closely. Each of the non-compliers (n = 149) was asked the following question: "If you were to improve the conditions under which your firearms are stored, what would be your first step?" (A list of options was read out, but the respondent was asked to choose only one). The two options most frequently mentioned by non-compliers were to render the firearm inoperable, that is equip it with a secure locking device or remove a part needed to discharge the weapon (40%; n = 60); and to ensure that the weapon is inaccessible, in other words, store it in a place that is locked and cannot readily be broken into (24%; n = 36).

The non-compliers were then asked the following questions: "Why is your firearm stored in an operable condition and in a place where it is accessible? What would prompt you to render your weapon both inoperable and inaccessible?" (No answers to this

FIGURE 1
Distribution of non-compliers (n = 175/504) according to firearm (shotgun or rifle) storage criteria studied



question were suggested and only one answer was recorded for each non-complier). In each case, the storage condition at issue was clearly defined to ensure that the respondent fully understood the questions (operable/inoperable; accessible/inaccessible).

Operable firearm

In explaining why their long guns were operable, the 149 non-compliers stated that they had taken other safety precautions (19%; n = 28) and that their weapon was safely hidden away (7%; n = 10). Negligence was also invoked as an explanation by a number of non-compliers (7%; n = 10). It should be noted that 15% (n = 22) of non-compliers could offer no particular reason to explain why their weapons were stored at home in an operable condition. The presence of children was most frequently invoked by non-compliers as a reason that would prompt them to render their long guns inoperable (28%; n = 42). An equal percentage of respondents (28%; n = 42) saw no particular reason to render their firearms inoperable.

Accessible firearms

A significant percentage of the 149 non-compliers indicated that they left their firearms in an accessible place either through negligence or force of habit (13%; n = 19), or because they felt the firearms were well hidden (9%; n = 13). Not having any other suitable place to store a gun was also mentioned by a number of non-compliers (8%; n = 12). It is important to note that 15% (n = 22) of non-compliers could offer no particular reason why their firearms were left in an accessible place. Again, the presence of children was most frequently invoked by non-compliers as a reason that would prompt them to make their long guns inaccessible (26%; n = 39). However, a significant percentage of non-compliers (35 %; n = 52) saw no particular reason to make their firearm inaccessible.

Discussion

Strengths and limitations

Several aspects of this survey must be emphasized:

- the participation rate was 65%, which is satisfactory given the subject matter (storage of firearms) and the data collection method (telephone survey);
- the participants were drawn from a random sample of 4,654 households, the distribution of which was based on the demographic weight of the various administrative regions of Quebec, which is likely to ensure more representative results;
- the state of long gun storage was described with respect to federal regulations on firearm storage, which not only constituted a first, but also provided a basic measure of the level of compliance with these regulations in Quebec;
- data on firearm storage came from the owners themselves rather than a third party, which ensured greater validity;[15,16]
- each participant was questioned on the storage of a single long gun at a specific point in time (the time of the interview), which tended to reduce the kind of information bias that can result from relying on memory;
- the final judgment on compliance with firearm storage regulations was made by the researchers rather than the gun owners, which was an advantage given the relative complexity of this type of judgment (where three criteria must be considered).

The results of this survey on long gun storage are subject to three types of bias. The first is that firearm storage practices constitute a form of reported behaviour. Some participants may deliberately have indicated that their firearms were stored more securely than they in fact were, which would lead the researchers to underestimate the number of non-compliers.

The second bias is linked to the voluntary aspect of participation in the survey. It may be that those who chose to participate were, on average, more likely to comply with firearm storage regulations than those who refused to take part, which would also cause the researchers to underestimate the number of non-compliers.

The third bias concerns the fact that participants were questioned about a single long gun. In cases where several firearms were being stored at home, it is possible that the firearm selected as the subject of the questionnaire may have been stored under different conditions than the other weapons. If storage regulations were being adhered to solely in the case of the selected firearm, this too would lead researchers to underestimate the number of non-compliers; if the opposite were true, there would be no bias.

Conditions of storage

Based on the results of this survey, at least 35% of persons who store one or more long gun at home fail to comply with Canadian firearm storage regulations.[13] It should be remembered that 89% (n = 156) of non-compliers stated that they were not the sole occupants of their household (in 59% of cases the other occupants were children < 18 years old), and 29% (n = 51) stated that their gun was accessible to others, usually a spouse (70%) or a child (40%).

The results of the survey lead us to estimate that at least 6% of Quebec homes contain at least one long gun that is improperly stored. This estimate was calculated by multiplying the percentage of Quebec households where at least one firearm is stored by an adult (17%) by the percentage of non-compliers among study subjects (35%).

The study did not link storage practices with firearms training as the same number of compliers and non-compliers had been trained. This is likely due to the fact that most of the participants had never received adequate instruction on the storage of firearms. In fact, at the time of the survey (1994), 91% of participants indicated having received training in the use of firearms in the past. However, before the Canadian firearm storage regulations came into effect (January 1993), the safe storage of long guns was not part of the firearms training courses. When the regulations came into effect, this deficiency was remedied, but the "enriched" training was only required for new gun owners, which likely affected only a small proportion of the study participants since it was carried out in 1994.

To our knowledge, there is no basis for comparison of firearm storage practices in Quebec and elsewhere. However, a survey conducted in 1999 with 282 long gun owners from all 10 provinces is somewhat interesting.[17] Methodological differences, however, make it rather risky to draw comparisons between the two surveys (Appendix 1). The study populations are not truly comparable, the criteria used to evaluate long gun storage conditions are considerably different, and the survey questions did not relate to the same number of weapons.

The main results of the 1999 survey are nonetheless presented by way of indication: 99% of the participants in that study stated that all the long guns they kept at home were stored unloaded; 83% indicated that all the long guns they kept at home were either kept under lock and key or had been rendered inoperable; and 98% of participants indicated that the ammunition for their firearms was stored securely, either in a separate place, or with the weapon, but in a locked compartment. Approximately 17% of participants had failed to comply with at least one storage criterion. Interestingly, similar results are observed in the 1994 survey if the 1999 storage criteria are applied (results not presented).

Courses of action

The study findings suggest some possible methods of increasing the level of compliance with Canadian firearm storage regulations.[13] In 85% of cases where the owner of a long gun had failed to comply with federal regulations (149/175), the weapon in question was both operable and accessible. To increase the level of compliance, an important goal should be to encourage the owners of long guns to render them inoperable (e.g. by use of a locking device) or inaccessible (e.g. by storing them in a securely locked place). These two measures are the ones most frequently identified by the non-compliers as a way of improving security. The survey findings also show that to persuade non-compliers to take these measures, they must be made aware of the fact that they will be protecting their children, as well as the children of neighbours and relations.

Study results indicate that reducing the number of firearms stored at home is another possible solution. More than a third (37%) of participants indicated that their gun(s) had not been used by themselves or by anyone else during the 12-month period preceding the survey. It should be noted that these percentages were even higher among non-compliers than compliers. It would seem important, therefore, to encourage those who own guns that they do not use to dispose of these weapons.

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We wish to thank Michel Lemieux and Pierre Duchesnay of the Léger & Léger polling firm in Quebec City who oversaw the data collection. This survey would not have been possible without the support of the Comité intersectoriel sur la violence familiale et l'entreposage sécuritaire des armes à feu au Québec (CCAAF), the Canadian Firearms Centre at the Department of Justice, and the regional public health branches of Quebec City and Montérégie.

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APPENDIX
Methodological differences between the two surveys: Quebec (1994) and Canada (1999)

***Populations not comparable*:** In the 1999 Canadian survey, only 40 participants out of 282 were from Quebec. This is significant in that in the 1994 survey, firearm storage conditions were associated with the language spoken by participants.

***Nature of the storage criteria*:** Contrary to the 1994 survey, the 1999 survey defines a location as being inaccessible if it is kept under lock and key, regardless of whether it is difficult to break open or into. As a result of these less stringent requirements with respect to accessibility, a greater number of participants were classified as being in compliance with two of the three criteria stipulated in federal firearm storage regulations, namely the criterion that requires long guns to be stored in such a way as to render them inoperable or inaccessible (second criterion) and the criterion stating that a weapon can be stored with its ammunition provided that they are stored in an inaccessible compartment or receptacle (third criterion).

***Number of long guns considered*:** In 1994, compliance with the storage criteria was evaluated by questioning each participant about a single long gun (when more than one firearm was stored at home, one of them was selected at random), whereas in the 1999 study, participants were asked to consider the storage conditions for all of the firearms in their possession.

Book Review

Design and Analysis of Cluster Randomization Trials in Health Research

By Allan Donner and Neil Klar
London (England): Arnold Publishers, 2000;
x + 151 pp; ISBN 0 340 69153 0; (hardcover)

The increasing popularity of the cluster randomization design among health researchers over the past two decades has led to an extensive body of methodology and a growing literature that cuts across several disciplines in the statistical, social and medical sciences. This book is the first to provide a unified and systematic treatment of the topic. It may be used as a reference source for investigators in the planning or analysis stages of a study or as a textbook for a graduate level course in research methodology.

The book includes fairly non-technical chapters summarizing key issues of study design, data analysis and reporting as well as more technical material describing extensions of standard regression models (e.g. generalized estimating equations approach, multilevel models) that are needed to account for the variance inflation due to clustering.

The book also provides intriguing discussions of the historical development of cluster randomized trials and summarizes the unique ethical challenges of cluster randomization. For example, in community randomized trials it is typically not possible to obtain informed consent from all individuals who may be affected by the intervention prior to random assignment. This would be the case, for example, in trials evaluating innovative methods of water treatment for the prevention of infectious diseases or in trials evaluating smoking cessation interventions using mass media.

This is a well-written book, which includes data and worked examples illustrating methods of sample size estimation and data analysis.

A challenge facing the authors is that the methodology for cluster randomization trials is undergoing very rapid development. For instance, since publication of the book, special issues of two leading journals[1,2] have been devoted to cluster randomization. It is to be hoped that a second edition is being considered in which the authors can discuss some of these newer developments. For example, as the number of trials adopting cluster randomization has increased, meta-analyses of trials using various units of allocation are starting to appear in the literature. However, investigators have very little guidance, as yet, on how best to conduct such meta-analyses.

Purchasers of the book are entitled to a 25% discount on ACluster, a computer software compatible with Windows 3.1, 95, 98 and NT that implements many of the sample size and analysis formulas presented in the book. Information concerning the software can be obtained at <http://www.arnoldpublishers.com/support/cluster/>.

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New Resource

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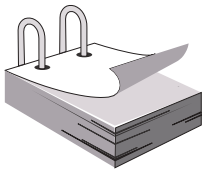

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May 28–30, 2001 **Toronto, Ontario**	Second International Symposium on the Effectiveness of Health Promotion Centre for Health Promotion University of Toronto	<www.utoronto.ca/chp>
June 13–16, 2001 **Toronto, Ontario**	Congress of Epidemiology 2001 Combined Meeting of American College of Epidemiology, American Public Health Association's Epidemiology Section, Canadian Society for Epidemiology and Biostatistics and Society for Epidemiologic Research	<www.epi2001.org>
July 1–6, 2001 **Vancouver, British Columbia**	"Global Aging: Working Together in a Changing World" 17th Congress of the International Association of Gerontology	Congress Secretariat Gerontology Research Centre Simon Fraser University 2800 – 515 West Hastings Street Vancouver, BC V6B 5K3 Tel.: (604) 291-5062 Fax: (604) 291-5066 E-mail: gero@sfu.ca <www.harbour.sfu.ca/iag>
July 15–20, 2001 **Paris, France**	"Health: An Investment For a Just Society" XVII World Conference on Health Promotion and Health Education International Union for Health Promotion and Education	Martine Lapergue Réjane Jouan Comite francais d'Education pour la Sante (C.F.E.S) XVII World Conference on Health Promotion and Health Education 2 , rue Auguste Comte - 92174 VANVES Cedex - FRANCE Tel.: 33 (0)1 41 09 96 48 Fax: 33 (0) 1 46 45 00 45 E-mail: mlapergue.cfes@imaginet.fr <www.iuhpe.org>
September 4–7, 2001 **Atlanta, Georgia, USA**	"Using Science to Build Comprehensive Cancer Programs: A 2001 Odyssey" US Department of Health and Human Services Centers for Disease Control and Prevention 2001 Cancer Conference	Conference Registration: June 1, 2001 American Cancer Society, National Home Office Attn: CDC's 2001 Cancer Conference 1599 Clifton Road, NE Atlanta, Georgia 30329 USA
September 22–25, 2001 **Sydney, Australia**	4th International Conference on the Scientific Basis of Health Services Sydney Convention and Exhibition Centre – Darling Harbour	<www.icsbhs.org>

October 18–21, 2001 **Saskatoon, Saskatchewan**	"Health Research in Rural and Remote Canada: Taking the Next Steps" Second Scientific Conference and Annual Meeting of The Rural Health Research Consortium	The Rural Health Research Consortium 103 Hospital Drive PO Box 120, RUH Saskatoon, SK S7N 0W8 Tel.: (306) 966-7888 Fax: (306) 966-8378 E-mail: <www.usask.ca/medicine/agmedicine>
November 29–December 1, 2001 **Toronto, Ontario**	Canada's Fifth National Conference on Asthma and Education (ASED 5) *Abstract Deadline: June 30, 2001*	A. Les McDonald, Executive Director Canadian Network for Asthma Care (CNAC) 1607 – 6 Forest Laneway North York, Ontario M2N 5X9 Tel.: (416) 224-9221 Fax: (416) 224-9220 E-mail: <www.cnac.net>

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