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ORIGINAL ARTICLE

# Examining the association between socioeconomic position and body mass index in 1978 and 2005 among Canadian working-age women and men

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**Abstract**
*Objectives* We examined the association between socioeconomic position (SEP) and body mass index (BMI) among Canadian men and women in 1978 and 2005. We examined both the average SEP–BMI association, and variation in this association across the distribution of BMI.
*Methods* We analysed data from two nationally representative surveys containing measured height and weight data: the Canada Health Survey (1978) and the Canadian Community Health Survey (2005). Ordinary least squares and quantile regression were used to examine average and distributional SEP–BMI associations, respectively, for each survey.
*Results* Education was inversely associated with BMI for men and women at both time points, and there was no evidence of narrowing between 1978 and 2005. This association was stronger for women than men, and was particularly strong for heavier women. Education and income related differently to BMI.
*Conclusions* The SEP–BMI association in Canada is complex, showing variation by gender, by aspect of SEP, across the BMI distribution, and at different time points. The association departs from the more consistent social gradient in health, thereby challenging our view of BMI as a typical health issue.

**Keywords** Body mass index (BMI) · Socioeconomic position (SEP) · Canada · Population trends

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## Introduction

There is a large literature on the association between socioeconomic position (SEP, referring to one's position in society as influenced by social and economic factors, Shaw et al. 2008) and obesity, body mass index (BMI), and other indicators of adiposity (McLaren 2007; Ball and Crawford 2005; Mikolajczyk and Richter 2008). The association shows gender differences, being predominantly inverse among women living in industrialised nations (i.e., higher SEP, lower BMI), while it is less consistent among men (McLaren 2007; Ball et al. 2002). Further variation is evident when aspects of SEP are considered. In Canada for example, data suggest an inverse association between education and BMI for both men and women, while the income–BMI association tends towards positive for men (Tjepkema 2006; McLaren and Godley 2009; Auld and Powell 2006). In this way, the socioeconomic patterning of weight departs from the familiar social gradient in health, whereby lower SEP is associated with poorer health status across an array of health outcomes (Mackenbach et al. 2008; Raphael 2009).

From a surveillance perspective, it is important to examine change over time in the association between SEP

and BMI since evidence of narrowing or widening disparities can illuminate the impact of changes in the social, economic, or political context. Trends over time in the socioeconomic patterning of weight have been examined in other countries (Gutiérrez-Fisac et al. 2002; Wardle and Boniface 2008; Wolff et al. 2006). For example, using population-based data (NHANES) from 1971 and 2000, Zhang and Wang (2004) observed that educational disparities in obesity in the United States appear to have narrowed over that time period, especially for women.

Canadian data on socioeconomic trends in weight are sparse and suffer from limitations such as self-reported height and weight data in some national surveys (e.g., Millar and Stephens 1993), which are known to be biased towards underestimating BMI, particularly among heavier individuals (Shields et al. 2008). Torrance et al. (2002) presented trends in the prevalence of obesity by education group using three nationally representative datasets (1970/1972, 1978, and 1986–1992) that contain measured height and weight data. Among men, the largest increase in obesity prevalence during this time period was observed in those with lowest education, while in women, trends were more uniform across the different education groups, with the largest obesity increase in the middle category.

To complement research focusing on categories of BMI (e.g., overweight, obesity), in this analysis we treat BMI as a continuous variable. Studies from Canada (Torrance et al. 2002; Katzmarzyk 2002), the US (Flegal and Troiano 2000), and the UK (Wardle and Boniface 2008) have shown that, over recent decades, mean BMI has increased, but the changes in the distribution of BMI are more complex (i.e., increasing skewness) such that they are not well-approximated as a location shift. The methods we use here can flexibly model such complex changes in the relationship between SEP and BMI. The nature of these changes in Canada is not currently known.

Our objective was to examine the association between SEP and BMI (average and distribution) among Canadian men and women at two points in time. To achieve this objective, we followed three procedural steps: first we assessed the average SEP–BMI association, adjusting for age, in 1978 and 2005. Second, we examined the SEP–BMI association across the BMI distribution, adjusting for age, at the two time points. Third, we explored the nature of the distributional shift over time by comparing the SEP–BMI quantile distributions in 1978 and 2005. We focus on Canada, which, despite having universal health care and a historically egalitarian reputation (http://hdr.undp.org/en/statistics/), has shown increasing socioeconomic inequality in recent years (Curry-Stevens 2009), as well as increasing obesity prevalence among both adults and children (Tjepkema 2006; Shields 2006) and is thus an interesting case study.

## Methods

### Design and data source

This study has a repeated cross-sectional design (sometimes called a trend design), wherein we have comparable data from two time points, with a different sample each time (Bryman 2004; Bowling 2002). We analyse data from two nationally representative Canadian health surveys (1978 Canada Health Survey; 2005 Canadian Community Health Survey; public use versions). Details for the surveys are available at http://www.statcan.gc.ca and in the report by Health and Welfare Canada (1981). Briefly, the objective of both surveys was to provide information on the health status of the Canadian population for research and planning purposes. The target population for both surveys was the non-institutionalised Canadian population living in the ten provinces (1978 and 2005) and three territories (2005 only); with certain exclusions such as those living on Indian Reservations and in remote regions. A stratified cluster sampling approach was used in both cases, and national response rates of 79% for the 2005 CCHS, and 72–89% (range for different survey components) for the 1978 CHS were achieved. Data collection occurred through a combination of methods (in person interview; self-completion questionnaire; telephone interview; physical measurement) facilitated by trained staff.

We used the following variables: BMI (kg/m$^2$, computed from height and weight which were measured by trained interviewers using equipment routinely checked for reliability), respondent education (two categories: bachelor's degree or higher, vs. less than complete bachelor's degree), personal income from all sources (three groups), sex, and age. For total personal income we began with the available variables, namely, a derived five-category variable in 2005 and a semi-continuous (grouped to nearest $500) variable in 1978. We used the Bank of Canada's online inflation calculator to create a five-category variable in 1978 that was comparable (in adjusted dollar value) to the existing 2005 variable. Due to the small size of some of the categories, we collapsed the five categories to three (higher, medium, and lower) that were as balanced in size as possible. The three categories were created separately for men and women to reflect the different income distributions (i.e., women's income is lower than men's across the distribution, particularly in 1978). The resulting income groups for men were: lowest [$0–$10,000 in 1978 ($n$ = 259) and $0–$29,999 in 2005 ($n$ = 298)]; middle [$10,500–$17,000 in 1978 ($n$ = 331) and $30,000–$49,999 in 2005 ($n$ = 344)]; and highest [$17,500–$30,000+ in 1978 ($n$ = 366) and $50,000–$80,000+ in 2005 ($n$ = 438)]. Income groups for women were: lowest [$0–$5,000 in 1978 ($n$ = 801) and $0–$14,999 in 2005 ($n$ = 383)];

middle [$5,500–$10,000 in 1978 ($n = 242$) and $15,000–$29,999 in 2005 ($n = 324$)]; and highest [$10,500–$30,000+ in 1978 ($n = 198$) and $30,000–$80,000+ in 2005 ($n = 492$)].

In the 1978 dataset, extreme values of height and weight were grouped using the third and 97th percentiles as a guide (Canada Health Survey Data Users' Guide). This was done to preserve confidentiality of respondents in the public use version of the data, which is the only version available (personal communication, Statistics Canada, 11 March 2009). To enable comparison between 1978 and 2005, we performed a similar truncation procedure in both datasets, whereby values representing the 3rd percentile and below, and the 97th percentile and above, were removed. In the event where an effect differed between the truncated and non-truncated versions of the 2005 data, this is noted.

## Analysis

Our analytic strategy was threefold, corresponding to the three procedural steps noted above. Analyses included appropriate survey weights to account for the complex survey sample design. All analyses were conducted using Stata version 10.

First, we examined the average SEP–BMI association, using ordinary least squares (OLS) regression, in 1978 and 2005. We regressed BMI onto education and income in the same model (dummy variables, lowest category as reference) for each survey separately, adjusting for age. Men and women were analysed separately.

Second, we used quantile regression (Hao and Naiman 2007; Koenker and Hallock 2001) to examine the SEP–BMI association across the BMI distribution, in 1978 and 2005, for men and women. We describe quantile regression by contrasting it with OLS regression. OLS regression models the conditional mean of BMI, as follows (abstracting from the categorical nature of the covariates):

$$E(\text{BMI}_i|\ \text{income}_i,\ \text{education}_i,\ \text{age}_i) = \beta_0 + \beta_1(\text{income}_i) + \beta_2(\text{education}_i) + \beta_3(\text{age}_i) \quad (1)$$

where $E$ denotes the expectation operator and $\beta$ are parameters to be estimated. An estimate of $\beta_1$ recovers how much BMI changes *on average* when we consider a respondent with an extra dollar of income, holding education and age constant. OLS estimates do not reveal changes in the distribution of BMI other than changes in the conditional mean. For example, a mean-preserving spread (i.e., a change such that the mean of BMI does not change, but more people are found at both the upper and lower extremes of weight) will not be recovered by OLS estimation. Since we are interested in the distribution of BMI (above and beyond the average), we overcome this limitation of OLS, following previous research (e.g., Kamhon and Wei-Der 2004; Quintana-Domeque 2005; Auld and Powell 2009), by turning to quantile regression models (also known as Least Absolute Deviation models) of BMI. We estimate models of the form:

$$q_\theta(\text{BMI}_i|\ \text{income}_i, \text{education}_i, \text{age}_i) = \gamma_{(\theta,0)} + \gamma_{(\theta,1)}(\text{income}_i) + \gamma_{(\theta,2)}(\text{education}_i) + \gamma_{(\theta,3)}(\text{age}_i) \quad (2)$$

where $\theta$ is a quantile (percentile) of the distribution of BMI and the $\gamma$ are parameters to be estimated. For example, at $\theta = 50$, the model is a median regression, and $\gamma_{50,1}$ recovers how much the median BMI changes as consider a respondent with an extra dollar of income, holding age and education constant. Similarly, at $\theta = 90$, the model recovers the determinants of the 90th percentile of BMI, and so on. We estimate the model at quantiles $\theta = 2$ through $\theta = 98$ in steps of two, stratifying by survey wave, such that the complete set of estimates consists of 50 estimated sets of parameters for each wave.

One drawback of quantile regression is the large amount of output, which can be somewhat unwieldy to present. Thus, a convenient manner of expressing these estimates compactly is to graph the estimated coefficients against quantiles, as we do in Figs. 1, 2, 3 and 4. In such figures, if the estimated coefficients do not vary across quantiles, then the data suggest that a one-unit change in the covariate induces a shift in the distribution of BMI, preserving the shape of the distribution. If the coefficients rise or fall as we consider higher quantiles, then the distribution of BMI changes with the covariate in a manner more complicated than a level shift. We implement these models using the "qreg" command in Stata 10, which uses a linear programming algorithm aimed at minimising the sum of absolute weighted residuals to solve for the estimated regression coefficients (in contrast to OLS which minimises the sum of squared residuals). In quantile regression, the residuals at each quantile are weighted using a multiplier of $2q$ ($q$ = quantile) if the residual is positive, and $2(1 - q)$ otherwise (StataCorp 2007). For income and education categories, we test the hypothesis that the estimated quantile regression coefficients are equal in 1978 and 2005 at each of the 50 quantiles between 2 and 98 at which we estimate the models. Since the samples are independent, the estimated coefficients are also independent, and asymptotically normally distributed under standard regularity conditions (see for example, Hao and Naiman 2007). The statistic:

$$\frac{\gamma_{\theta j,1978} - \gamma_{\theta j,2005}}{\sqrt{V(\gamma_{\theta j,1978}) + V(\gamma_{\theta j,2005})}} \quad (3)$$

(where all parameters are estimated values, and $\gamma_j$ is the $j$th estimated parameter) is then asymptotically standard normal under the null that the coefficients at quantile $\theta$ are

equal in 1978 and 2005. Since we conduct 50 such tests for categories of education and income, we present the results using bold vertical lines to indicate locations on the figures where we reject the null (that estimated coefficients are equal in 1978 and 2005) at 5% nominal size.

## Results

Of the 8,486 individuals in the pertinent 1978 subsample (for whom height and weight data were potentially available), 4,005 were age 25–64 years, and height and weight data were available for 2,869 (71.6%) of these. Those with missing BMI data were more likely to be male (58.1% vs. 41.9% female) and younger (56.1% were age 25–44 vs. 43.3% age 45–64). Of the 4,735 respondents in the pertinent 2005 subsample, 2,769 were age 25–64, and height and weight data were available for 2,705 (97.7%) of these.

Missing data on education occurred in <1% of cases, and missing data on income ranged from a low of 5.2% for men in 2005 to a high of 14.7% for men in 1978 (the amount of missing data for women fell between these extremes). There was one instance of a significant difference at $P < 0.05$ between those with missing versus non-missing SEP data on BMI: in 1978, men with missing income data were heavier than men with non-missing income data (mean BMI 26.2 vs. 25.6 kg/m$^2$, $P = 0.048$).

Analyses are based on those with complete data on all variables of interest (BMI, sex, age, income, and education): $n = 2{,}197$ (956 men and 1,241 women) in 1978, and $n = 2{,}279$ (1,080 men and 1,199 women) in 2005. Descriptive statistics for both survey samples are presented in Table 1.

Results of OLS and quantile regression are presented in Table 2 and Figs. 1, 2, 3 and 4, respectively. All results adjust for age, with 10-year age groups entered as dummy variables.

**Table 1** Descriptive statistics for study samples, Canadian adults aged 25–64 years, 1978 and 2005 (weighted estimates)

| | Men<br>BMI (kg/m$^2$): mean (SD), *n* (% of total) | Women<br>BMI (kg/m$^2$): mean (SD), *n* (% of total) |
|---|---|---|
| Canada Health Survey, 1978 | | |
| Full sample | 25.7 (3.3), $n = 956$ (range 16.9–37.4) | 24.6 (4.1), $n = 1{,}241$ (range 16.3–37.7) |
| Education | | |
| <Univ degree | 25.8 (3.4), $n = 826$ (86.4%) | 24.8 (4.1), $n = 1{,}148$ (92.5%) |
| Univ degree+ | 24.9 (2.7)*, $n = 130$ (13.6%) | 22.7 (2.8)*, $n = 93$ (7.5%) |
| Total (education) | 956 | 1,241 |
| Personal income[A,B] | | |
| Lowest | 25.7 (3.6)[a], $n = 259$ (27.1%) | 25.2 (4.2)[a], $n = 801$ (64.5%) |
| Middle | 25.7 (3.6)[a], $n = 331$ (34.6%) | 24.1 (3.9)[b], $n = 242$ (19.5%) |
| Highest | 25.7 (3.0)[a], $n = 366$ (38.3%) | 23.4 (3.4)[b], $n = 198$ (16.0%) |
| Total *n* (income) | 956 | 1,241 |
| Canadian Community Health Survey (Cycle 3.1), 2005 | | |
| Full sample | 27.6 (3.8), $n = 1{,}080$ (range 18.2–41.2) | 26.3 (5.1), $n = 1{,}199$ (range 16.0–45.7) |
| Education | | |
| <Univ degree | 28.0 (3.7), $n = 384$ (35.6%) | 27.4 (5.4), $n = 432$ (36.0%) |
| Univ degree+ | 27.3 (3.8)*, $n = 696$ (64.4%) | 25.8 (4.9)*, $n = 767$ (64.0%) |
| Total *n* (education) | 1,080 | 1,199 |
| Personal income[A,B] | | |
| Lowest | 27.8 (4.0)[a], $n = 298$ (27.6%) | 26.1 (5.2)[a], $n = 383$ (32.0%) |
| Middle | 27.1 (3.7)[a], $n = 344$ (31.9%) | 26.9 (5.3)[a], $n = 324$ (27.0%) |
| Highest | 27.8 (3.7)[a], $n = 438$ (40.6%) | 26.2 (4.9)[a], $n = 492$ (41.0%) |
| Total (income) | 1,080 | 1,199 |

[A] Superscripts should be read across income categories, and different superscripts indicate significant differences in BMI at $P < 0.05$ (based on ANOVA with post hoc tests)

[B] Income categories were created separately for men and women, to reflect sample distribution

* Differs from <Univ degree at $P < 0.05$

**Table 2** Results of OLS regression analysis examining the association between socioeconomic position and BMI, conducted separately for men and women, 1978 and 2005 (weighted estimates, adjusting for age)

| Predictor variables | Canada Health Survey, 1978 | | | |
|---|---|---|---|---|
| | Model A: men ($n = 956$) | | Model B: women ($n = 1{,}241$) | |
| | Coefficient (SE) | 95% CI | Coefficient (SE) | 95% CI |
| Educ: Univ degree+ | −0.78 (0.31)* | −1.4 to −0.17 | −0.96 (0.42)* | −1.8 to −0.14 |
| Income: middle | 0.11 (0.30) | −0.48 to 0.69 | −0.69 (0.28)* | −1.2 to −0.14 |
| Income: highest | 0.13 (0.30) | −0.47 to 0.72 | −0.91 (0.31)** | −1.5 to −0.30 |
| Constant | 25.4 (0.28) | | 23.4 (0.21) | |
| Predictor variables | Canadian Community Health Survey (Cycle 3.1), 2005 | | | |
| | Model C: men ($n = 1{,}080$) | | Model D: women ($n = 1{,}199$) | |
| | Coefficient (SE) | 95% CI | Coefficient (SE) | 95% CI |
| Educ: Univ degree+ | −0.57 (0.24)* | −1.05 to −0.09 | −1.3 (0.32)** | −1.9 to −0.67 |
| Income: middle | −0.30 (0.30) | −0.89 to 0.30 | 0.85 (0.38)* | 0.10–1.6 |
| Income: highest | 0.27 (0.29) | −0.31 to 0.84 | 0.36 (0.35) | −0.31 to 1.04 |
| Constant | 26.8 (0.32) | | 25.3 (0.43) | |

Reference category is lowest (less than complete bachelor's degree; lowest income group)

Income categories were created separately for men and women, to reflect sample distribution

Ten-year age groups (i.e., 25–34, 35–44, 45–54, and 55–64 years), entered as dummy variables

* $P < 0.05$, ** $P < 0.01$

For women, OLS regressions for both 1978 and 2005 (Table 2, Models B and D) indicate that higher education was associated with lower mean BMI, holding age and income constant. Looking at Fig. 1, the education–BMI association varies across the BMI distribution in both 1978 (dashed line) and 2005 (solid line): in both cases it is more strongly inverse (departs more from zero in a negative direction) for women at the higher (heavier) end of the BMI distribution. The education coefficient for women differed significantly (at $P < 0.05$) from the base category at the following BMI quantiles: 62, 76–86, and 90–96 in 1978; and 8–18, 24, 54–80, and 84–92 in 2005 (not indicated on figure). Bold vertical lines indicate the quantiles for which the 1978 and 2005 lines differ at $P < 0.05$. In this case, the single bold vertical line indicates minimal shift between the two time periods, which occurred at the lighter end of the BMI distribution, where the effect is more negative in 2005 than in 1978.

For men, as in women, OLS results (Table 2, Models A and C) indicate an inverse education–BMI association in both 1978 and 2005, holding age and income constant. In terms of sex differences (analyses not shown), we observed a sex × education interaction that was marginally significant ($P = 0.075$) in 1978 and significant ($P = 0.024$) in 2005, holding age and income constant. Figure 2 shows the education–BMI association across the BMI distribution for men. In 1978, the coefficients are below zero (negative effect) across the distribution, but particularly at the higher (heavier) end of the BMI distribution. This pattern is similar to, though less pronounced than, findings for women in 1978. For men in 1978, the education coefficient differed significantly (at $P < 0.05$) from the base category at quantiles 64–66 and 76–92 (not indicated on figure). In 2005, the negative coefficients appear more uniform across the BMI distribution, although they differed significantly from the base category at quantiles 6–10, 20–24, and 64–74 (not indicated on figure). The effect has become less negative at higher BMIs, as shown by the single bold vertical line in Fig. 2.

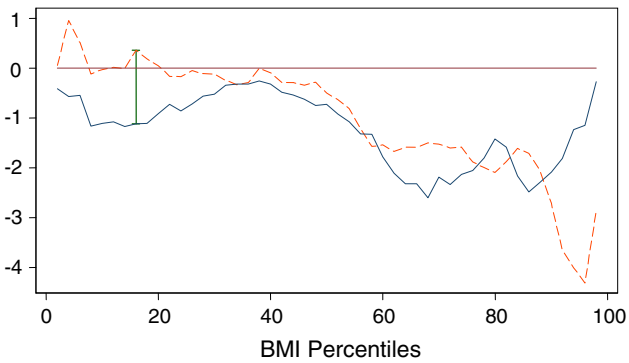


**Fig. 1** Results of quantile regression examining the effect (coefficient) of higher education (complete bachelor's degree) on BMI across the BMI distribution for women in 1978 (*dashed line*) and 2005 (*solid line*). *Dashed line* higher education, 1978. *Solid line* higher education, 2005. Base category is lower education (less than completed bachelor's degree). *Bold vertical line* denotes significant difference between the 1978 and 2005 lines at $P < 0.05$. Quantile regression model included education, income, and age

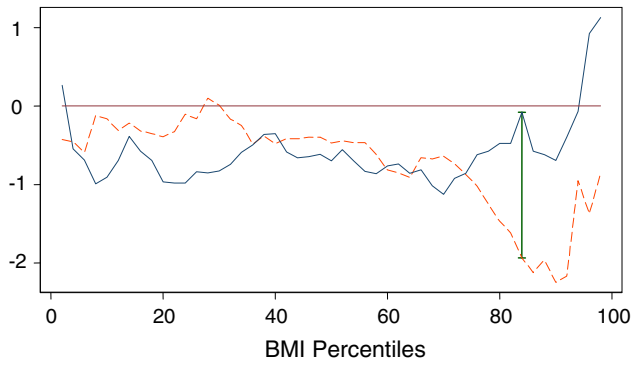


**Fig. 2** Results of quantile regression examining the effect (coefficient) of higher education (complete bachelor's degree) on BMI across the BMI distribution for men in 1978 (*dashed line*) and 2005 (*solid line*). *Dashed line* higher education, 1978. *Solid line* higher education, 2005. Base category is lower education (less than completed bachelor's degree). *Bold vertical line* denotes significant difference between the 1978 and 2005 lines at $P < 0.05$. Quantile regression model included education, income, and age

For the income–BMI association in women, OLS patterns showed an inverse gradient in 1978 (Table 2, Model B) whereby those of middle and highest income were lighter than those of lowest income, holding age and education constant. In 2005, in contrast, those of middle income were heavier than those of lowest income (Table 2, Model D). Post hoc tests of the survey × income interaction in women (not shown) confirm a significant change in OLS coefficient between 1978 and 2005, for both middle and high income ($P < 0.01$ in both cases). Looking at Fig. 3 (which features highest versus lowest income), it appears that in 1978, the negative effect for high income is particularly strong for those of higher BMI; whereas in 2005 the line hovers more uniformly above zero for most BMI quantiles (with some exception at the highest BMI

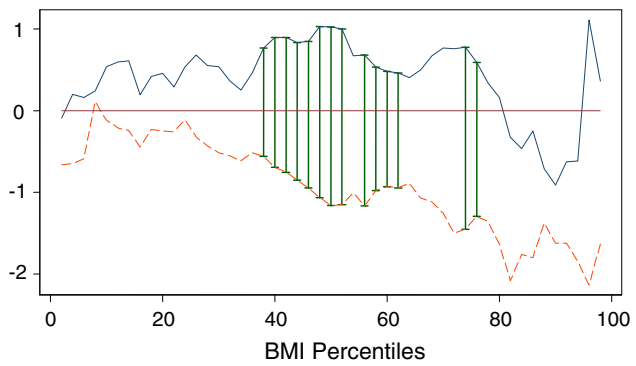


**Fig. 3** Results of quantile regression examining the effect (coefficient) of higher income on BMI across the BMI distribution for women in 1978 (*dashed line*) and 2005 (*solid line*). *Dashed line* higher income, 1978. *Solid line* higher income, 2005. Base category is lower income. *Bold vertical line* denotes significant difference between the 1978 and 2005 lines at $P < 0.05$. Quantile regression model included education, income, and age

values). The income coefficient for women differed significantly (at $P < 0.05$) from the base category at the following BMI quantiles: 40–52, 62, and 72–86 in 1978; and 10, 40–44, and 48 in 2005 (not indicated on figure). Shift between 1978 and 2005 was most prominent in the middle of the BMI distribution, where the effect of high income went from negative in 1978 to positive in 2005.

For men, OLS revealed no effect of income in 1978 (Table 2, Model A) or in 2005 (Table 2, Model C). Looking at the quantile effects in Fig. 4, the lines hover at or above zero for most BMI quantiles, in both 1978 (where the coefficient was observed to differ significantly ($P < 0.05$) from baseline only at quantiles 18 and 22) and 2005 [where the coefficient differed significantly ($P < 0.05$) from baseline at quantiles 6–10 and 66] (not indicated on figure). At the middle to higher end of the BMI distribution the line appears more positive in 2005 than in 1978 (supported by one quantile difference at $P < 0.05$), but the effect is not pronounced.

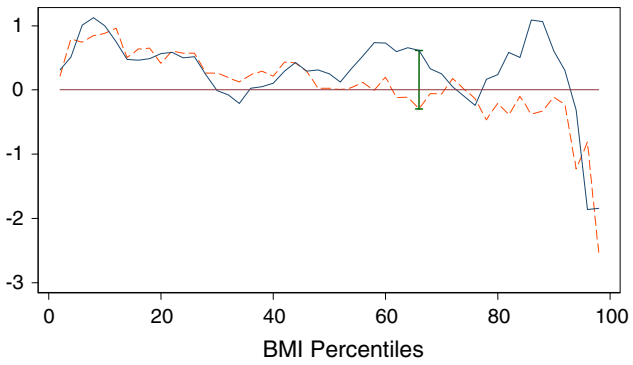


**Fig. 4** Results of quantile regression examining the effect (coefficient) of higher income on BMI across the BMI distribution for men in 1978 (*dashed line*) and 2005 (*solid line*). *Dashed line* higher income, 1978. *Solid line* higher income, 2005. Base category is lower income. *Bold vertical line* denotes significant difference between the 1978 and 2005 lines at $P < 0.05$. Quantile regression model included education, income, and age

We note that, using the non-truncated 2005 data, and when narrower income groups were examined, we observed (in OLS models) a significant ($P < 0.05$) positive effect of high income on BMI in men, which is consistent with other Canadian published data (Tjepkema 2006; McLaren and Godley 2009; Shields and Tjepkema 2006; Kuhle and Veugelers 2008).

## Discussion

Our main message is that the association between SEP and BMI among Canadian adults is complex: it varies by aspect of SEP, by gender, across the BMI distribution, and at

different time periods. The SEP–BMI association does not appear to follow the typical SEP-health gradient, which challenges a conceptualisation of BMI (including overweight, obesity) as a typical health issue.

Beginning with the association between education and BMI, OLS regression revealed an inverse association for both women and men, in both 1978 and 2005. The association was stronger for women, particularly in 2005, which accords with existing research (McLaren 2007). We found no evidence to suggest that educational inequalities in average BMI narrowed between 1978 and 2005. This finding contrasts with previous Canadian studies that suggested a decrease in educational disparities in obesity over time (Millar and Stephens 1993, Torrance et al. 2002); however, given that socioeconomic inequalities have increased in Canada since the late 1980s/early 1990s (Curry-Stevens 2009), we may have captured patterns that were not evident in these earlier papers. Our finding also contrasts with a study from the USA described in our Introduction section (Zhang and Wang 2004) which indicated a narrowing over time of educational disparities in obesity, particularly in women. This raises questions about possible Canada–USA differences; for example, perhaps some of the pertinent social and cultural drivers of obesity (e.g., social and structural norms and constraints around food, activity, and weight) are more pervasive and thus transcend educational categories to a greater extent in the US than in Canada. We confirmed in a post hoc analysis (not shown) that our findings are not an artefact of having focused on mean BMI as opposed to per cent obese.

Using quantile regression we examined the SEP–BMI association across the BMI distribution. For women, we observed a stronger inverse education–BMI association for heavier women in both 1978 and 2005. A similar pattern (though less pronounced) was observed for men in 1978, but not in 2005. When considering reasons for an inverse education–BMI association, one possibility is that higher education is associated with particular attributes that may facilitate a lower/healthier weight, such as health literacy, or affinity towards a health-promoting lifestyle. However, the observation that the effect was stronger for heavier women suggests that other things are at play. One likely factor pertains to social pressures around thinness and societal disparagement of fat (on aesthetic and health grounds), which are especially pronounced for women, and which are closely linked with socioeconomic hierarchy and aspirations of social distinction (Bourdieu 1984; Shilling 2005; Gortmaker et al. 1993; McLaren and Kuh 2004; Sargent and Blanchflower 1994). Because heavier women depart more from the thin ideal of physical beauty, a stronger class effect at higher BMI is not surprising. One question that arises when we consider recent increases in obesity prevalence is, have social pressures for thinness declined over this time period? Though indirect, our findings speak against this position: to the extent that these social pressures manifest as class differences in weight (reflecting stronger pressures in higher class groups), our findings suggest that social pressures remain and have perhaps intensified, especially for heavier women. For men, the education–BMI association was smaller, and it was more even across the BMI distribution in 1978 and particularly in 2005. This is consistent with weight-related social pressures being weaker for men, with gender differences especially prominent at higher BMIs.

Our findings with income differed from those with education, which speaks to the importance of examining these aspects of SEP separately. Although we observed no clear association between income and BMI for men in our data, we note that using the full (non-truncated) 2005 dataset we detected the positive effect (higher income, higher BMI) reported in other studies (Tjepkema 2006; McLaren and Godley 2009; Kuhle and Veugelers 2008), an effect for which various social and behavioural explanations are plausible (McLaren and Godley 2009). For women, the association between income and BMI changed between the two time periods. In 1978, it was inverse, and stronger for heavier women (similar to the pattern observed with education). In 2005, women in the middle income category were heavier on average than those in the lowest income category (this has been observed elsewhere, Tjepkema 2006), and the positive effect of income was fairly uniform across the BMI distribution. Insight into possible reasons for these findings may come from considering high earning women at the two time periods (since our analyses were based on personal income). In 1978, high earning women were fewer in number, and therefore more likely working in male-dominated environments. Perhaps the resulting gender-imbalanced setting conveyed workplace-based appearance-related norms that affected high income women of diverse body sizes, but particularly those of higher BMI. In 2005, it is conceivable that the sedentary nature of these jobs (which was likely also the case in 1978) coupled with other gendered aspects of contemporary lifestyle (e.g., child care, long commutes to work, other time pressures experienced by dual income families or single working mothers; McLaren et al. 2009) have contributed to higher BMIs and that these influences operate in a relatively non-discerning manner.

An important drawback of our study is the truncation of the 1978 data, which precluded examination of the extreme values of BMI (i.e., 3rd percentile and below; 97th percentile and above). This is an unfortunate consequence of secondary data analysis, particularly when the data are several decades old. To our knowledge there is only one published study that reports historical data on the income–BMI association in Canada (Shields and Tjepkema 2006), and we are pleased that our 1978 findings for women (inverse association between income and BMI) are consistent with that paper's

findings, which were based on data collected in 1986–1992. This lends some support to the validity of our 1978 findings, despite not having extreme values of BMI. Also, we acknowledge an additional limitation of both OLS and quantile regression, which is that these techniques recover associations but not causality.

Data limitations notwithstanding, we have shown that the relationship between SEP and BMI among Canadian adults is complex, varying by aspect of SEP, by gender, across the BMI distribution, and at different points in time. Strengths of our study include the nationally representative samples and the availability of data on measured height and weight, over a quarter of a century apart. Future studies might incorporate additional behavioural and sociodemographic variables towards understanding the mechanisms underlying the patterns observed here.

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