Is social assistance boosting the health of the poor? Results from Ontario and three countries

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Abstract

Intervention Social assistance programs supplement incomes of the most income-insecure. Because income is a fundamental source of health, income supplementation is expected to result in a boost to health status. As Canada finds itself in the midst of heated debate regarding the structuring (and restructuring) of social assistance programs, there is little evidence available for policymakers about the effectiveness of current social assistance programs in improving the health of the income-insecure.

Research question In this paper, we evaluate the health effects of social assistance programs in Ontario, Canada-wide and in peer programs from the United States and the United Kingdom.

Methods We used nationally representative household panel surveys (e.g., Canadian Survey of Labour and Income Dynamics) which follow individuals over time. Using fixed effects modelling, which controls for time-invariant characteristics of individuals, and further controlling for key time-varying characteristics, we modelled change in health status associated with change in receipt of social assistance in these societies. Health status was measured using self-rated health (fair/poor versus good/very good/excellent).

Results Our results suggest that the health of social assistance recipients was worse (Ontario, Canada, UK) or no different (US) than the health of non-recipients. For example, in Canada, receipt of social assistance was associated with 52.5% higher odds of reporting fair or poor health.

Conclusion Social assistance programs in Canada and peer countries are currently inadequate for improving the health of the income-insecure. This is likely due to insufficient benefits, exposure to precarious job conditions, or selection factors.

Résumé

Intervention Les programmes d'aide sociale complètent les revenus de la plupart des personnes en situation d'insécurité financière. Le revenu étant une source fondamentale de la santé, on peut s'attendre à ce que la supplémentation du revenu entraîne une amélioration de l'état de santé. Un débat sur la structuration (et la restructuration) des programmes d'aide sociale fait rage au Canada, mais les responsables des politiques ont peu de données probantes sur l'efficacité des programmes actuels pour améliorer la santé des personnes en situation d'insécurité financière.

Question de recherche Dans cet article, nous évaluons les effets sur la santé des programmes d'aide sociale en Ontario, dans l'ensemble du Canada, aux États-Unis et au Royaume-Uni.

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Méthode Nous avons fait appel à des enquêtes par panel représentatives à l'échelle nationale menées auprès des ménages (p. ex., l'Enquête sur la dynamique du travail et du revenu de Statistique Canada) qui suivent les répondants au fil du temps. À l'aide d'un modèle à effets fixes, pour tenir compte des caractéristiques des répondants qui sont invariables dans le temps, et après avoir apporté des ajustements pour tenir compte des effets des principales caractéristiques variables dans le temps, nous avons modélisé l'évolution de l'état de santé associée aux changements dans les prestations d'aide sociale dans les sociétés en question. L'état de santé a été mesuré selon un barème d'auto-évaluation de la santé (passable/mauvaise c. bonne/très bonne/excellente).

Résultats Nos résultats indiquent que la santé des bénéficiaires de l'aide sociale était moins bonne (Ontario, Canada, Royaume-Uni) ou la même (États-Unis) que celle des non-bénéficiaires. Au Canada par exemple, le fait de toucher des prestations d'aide sociale était associé à une probabilité 52,5 % plus élevée de déclarer une santé passable ou mauvaise.

Conclusion Les programmes d'aide sociale au Canada et dans des pays comparables ne suffisent pas actuellement à améliorer la santé des personnes en situation d'insécurité financière. C'est probablement en raison de l'insuffisance des prestations, de l'exposition à des conditions d'emploi précaires ou de facteurs de sélection.

Keywords Social assistance · Welfare · Social policy · OECD · Canada · United States · United Kingdom

Mots-clés Assistance sociale · Aide sociale · Politique sociale · OCDE · Canada · États-Unis · Royaume-Uni

Introduction

Health inequalities are widening in Canada and comparable nations (Hajizadeh et al. 2016). The poor have always been sicker and died sooner, but things appear to be getting worse. A primary hypothesis is that widening health inequalities reflect observed widening inequalities in income and other socio-economic resources (Siddiqi et al. 2013). This is because these resources are the 'fundamental causes' of health status (Link and Phelan 1995). They shape the everyday living conditions—experiences of stress, diet, physical activity, exposure to environmental toxins—which are the mechanisms that make us sick or keep us healthy.

Public health theory predicts that social policies, which reduce socio-economic inequalities, should be the most effective interventions for reducing health inequalities (Marmot 1999). In particular, income assistance programs, known in Canada as "Social Assistance (SA)", are used to supplement the incomes of the most income-insecure members of society during hard times of unemployment, or precarious employment, when earnings are insufficient to meet basic needs (by contrast, employment or unemployment insurance programs are designed to supplement incomes of the unemployed, irrespective of need). Therefore, an important public health research question is whether SA programs in Canada are effective, as theory would predict, in protecting the health of the income-insecure in society.

Currently, each Canadian province maintains its own SA programs, though there is significant consistency across many provinces, including Ontario, Alberta, and British Columbia. By and large, these are means-tested programs, which provide income and employment assistance to individuals whose households lack sufficient financial resources to meet a basic standard of living, as defined by their respective governments. Most recipients are expected to demonstrate reasonable effort

at seeking and accepting employment, else they risk losing their benefits.

Canada falls into a group of countries, the so-called liberal market economies, with others such as the United States and the United Kingdom, which are typified by means-tested social policies. The nature of these policies is to impose significant demands and provide limited benefits on the income-insecure (Esping-Andersen 1990). The US equivalent of SA is called Temporary Assistance for Needy Families (TANF), and is even more limited in eligibility and generosity than Canadian SA. TANF is available only to low-income households with children, has much more stringent work conditionalities, and lower benefit levels (Basu et al. 2016). In the UK, there are two separate benefit programs, Income Support (IS) and income-based Jobseeker's Allowance (JSA), which in combination are the functional equivalent of Canadian SA, with fewer conditionalities, and slightly more generous benefits (Farrants et al. 2016).

A review of the literature yielded approximately eight studies, which investigated the impact of social assistance on health outcomes among the liberal market economies, with only one based in Canada and the remainder based in the US (Basu et al. 2016; Vozoris and Tarasuk 2004; Dooley and Prause 2002; Ensminger and Juon 2001; Jayakody et al. 2000; Muennig et al. 2013; Narain et al. 2017; Wilde et al. 2014). In Canada, Vozoris and Tarasuk (2004) compared the odds of worse health among social recipients to that of non-recipients in a cross-sectional sample, finding worse health among social assistance recipients, after accounting for confounders such as age and education (Vozoris and Tarasuk 2004). Three US-based studies also conducted descriptive assessments and produced similar findings (Dooley and Prause 2002; Ensminger and Juon 2001; Jayakody et al. 2000). Four US studies used quasi-experimental designs which, compared to descriptive methods, are much more rigorous because they

are better able to isolate how much of the health difference between two groups can be attributed to a policy, rather than to other factors that make groups differ (Basu et al. 2016; Muennig et al. 2013; Narain et al. 2017; Wilde et al. 2014). These studies found a worsening of health among low-income Americans attributable to “welfare reform”, a restructuring of social assistance that took place during the mid 1990s, moving the US to the less-generous TANF model.

Thus, at this crucial moment in time, when health inequalities are growing, there are very few studies available—even fewer using rigorous methods from the perspective of causal inference—to understand whether income assistance programs are helping to buttress health of the income-insecure. The lack of evidence on Canadian SA is particularly glaring. At the same time, Canada and its peer countries are in the midst of hotly debating how SA programs should be delivered, and research like this can inform the debate by providing a sense of the health impact of current social assistance programs (CBC News n.d.).

In this study we pursue the question, do social assistance programs improve the health of society’s most income-insecure in Canada and peer liberal market economies? We do so by using a longitudinal, fixed effects study design, which, as we describe below, greatly improves upon traditional descriptive methodologies in terms of isolating the effects of a social policy from other, co-occurring factors. We examine Canada (as well as Ontario, as the largest province, and one undergoing considerable transition with respect to social assistance), the US, and the UK.

Methods

Data sources Nationally representative household panel surveys provide the best available data for investigating our hypothesis because: (a) they contain detailed information about employment status and collection of social assistance over time, (b) they permit establishment of temporal order between exposure and outcomes by exploiting the availability of data on change over time, and (c) they allow for control of time-invariant confounders.

For Ontario and Canada, we used the Survey of Labour and Income Dynamics (SLID), which contains several panels that start in varying years. We drew on the 2002–2007 and the 2005–2010 panels. For the US (Survey of Income and Program Participation, 2001–2003, 2004–2005, 2009–2011) and the UK (British Household Panel Survey, 2001–2008), we drew on panels most closely overlapped in time with SLID.

For all three datasets, we subsetting the general sample of each survey to an analytic sample through the following process. Eligible individuals were those who were present in the baseline year, which was the first year of their respective panel periods. That is, new entrants in subsequent years were excluded. Next, we included only those participants who had

experienced a change in their health status after the first year of the panel, so that we could measure change, and thus establish temporal order. As noted above, a key benefit of longitudinal analysis in terms of causal inference is the ability to measure change. We further restricted to those who, during the baseline year, were of working age (18–64 years old) and labour-force active. We eliminated individual observations if they had received Employment Insurance during the baseline year so that we could isolate the impact of SA from that of another income assistance program.

The proportion of missing values was less than 5% for any variable, and was thus addressed using listwise deletion. The final sample sizes for Ontario, Canada, the US, and the UK were 5367, 18,184, 98,556, and 7793 individuals (20,071, 69,058, 203,723, and 42,050 person-year observations), respectively.

The analytic samples are representative of the population groups that were included, but are not (necessarily) representative of the general population. Indeed, a routine trade-off between general descriptive methods, and more rigorous methods intended to provide tighter causal inference, is the representativeness of the sample. In order to assess any systematic differences, we compared the characteristics of the more general sample with our analytic sample (Table 1).

Outcome variable Health status was examined using self-rated health, which is generally the only metric of health available in household panel surveys, and is correlated with many other indicators of health status (e.g., chronic disease, depression, and mortality) (Franks et al. 2003). Self-rated health was reported by respondents on a five-category Likert scale and was dichotomized as fair/poor versus good/very good/excellent to produce adequate sample size in each analyzed category.

Receipt of social assistance benefits The main exposure variable was receipt of social assistance benefits by at least one household family member during the baseline year (the first year they were enrolled in the study). This is because the impact of income on an individual is not only a consequence of their personal income, but also may be a consequence of the income of the household in which they reside. For Ontario and Canada, it was derived from the question “Did [you] receive any income from social assistance or welfare in [baseline year]?” Those responding affirmatively were coded as exposed (1) and compared to those who did not report receiving any income from social assistance (0). For the US, the variable was coded exposed (1) if positive amount is reported anytime in the past 12 months for the variable asking “Total family public assistance payments such as Assistance for Families with Dependent Children (AFDC) or Temporary Assistance for Needy Families (TANF) for this month”, and compared to those who did not report any positive amount for any given month in the baseline year. For the UK, the BHPS asks “Have you yourself or jointly with others since [baseline year] received Income Support?”

Table 1 Weighted baseline characteristics: 18–64 in the labour force (no EI)

	Ontario		Canada		US		UK	
	(SLID 2002–2010)		(SLID 2002–2010)		(SIPP 2001–2011)		(BHPS 2001–2008)	
	Full <i>i</i> = 5367	Analytic <i>i</i> = 608	Full <i>i</i> = 18,184	Analytic <i>i</i> = 1983	Full <i>i</i> = 98,556	Analytic <i>i</i> = 6921	Full <i>i</i> = 7793	Analytic <i>i</i> = 3084
Age (years)	39.9	42.9	40.0	43.2	40.1	43.7	40.8	40.1
Female	47.5%	48.5%	47.4%	49.1%	47.20%	48.70%	44.8%	44.9%
Household type								
Single	22.7%	24.3%	23.4%	25.2%	23.4%	25.5%	11.7%	11.2%
Couple	18.2%	16.3%	21.5%	20.8%	19.5%	20.4%	44.3%	41.6%
Single with children	7.6%	9.7%	7.3%	8.0%	13.1%	17.0%	7.2%	7.7%
Couple with children	51.6%	49.8%	47.8%	46.0%	44.0%	37.1%	36.8%	39.6%
Have children	52.7%	56.6%	48.8%	50.6%	57.1%	36.0%	36.0%	38.8%
Number of children	1.00	1.04	0.92	0.90	1.09	1.01	0.63	0.69
Visible minority	21.7%	25.3%	17.9%	23.4%	17.5%	22.1%	3.5%	3.9%
Immigrant	28.2%	33.3%	19.9%	25.6%	15.3%	17.3%	5.3%	5.3%
Education								
Post-secondary	25.3%	15.4%	22.7%	17.4%	29.7%	18.5%	19.7%	17.9%
Some post-secondary	49.7%	54.8%	50.2%	49.2%	34.9%	34.4%	31.0%	31.1%
Secondary	14.4%	13.7%	15.3%	14.5%	26.5%	32.1%	34.6%	35.6%
Less than secondary	10.7%	16.1%	11.8%	18.9%	9.0%	15.1%	14.8%	15.4%
Home owner	76.6%	75.8%	74.2%	72.1%	70.7%	64.8%	83.6%	82.5%
Birth cohort								
Pre or 1940+	12.2%	14.4%	12.5%	17.2%	13.2%	19.0%	22.9%	19.8%
1950+	25.0%	33.4%	25.9%	32.1%	24.5%	31.7%	25.7%	26.6%
1960+	30.8%	34.5%	29.2%	31.9%	26.0%	25.0%	27.8%	30.3%
1970+	20.8%	12.3%	20.9%	12.5%	22.6%	16.5%	18.9%	18.4%
1980+	11.2%	5.4%	11.5%	6.3%	13.7%	7.8%	4.7%	4.9%
Annual household income	50,219	45,257	46,990	42,306	49,645	41,174	21,823	20,925
Health insurance coverage					80.2%	71.9%		
Regions ^a								
Region 1			6%	6%	18.2%	16.2%	4.2%	4.7%
Region 2			20%	22%	23.4%	23.0%	11.4%	11.7%
Region 3			24%	17%	36.2%	39.5%	9.1%	10.3%
Region 4			38%	40%	22.3%	21.3%	8.4%	9.6%
Region 5			13%	16%			8.5%	8.3%
Region 6							9.6%	8.6%
Region 7							9.3%	8.8%
Region 8							16.0%	15.5%
Region 9							7.9%	7.8%
Region 10							5.1%	5.0%
Region 11							9.1%	8.5%
Region 12							1.6%	1.2%
Social assistance recipient	5.2%	8.5%	5.10%	8.70%	11.5%	3.2%	0.0%	4.0%
Unemployed	4.0%	6.2%	3.80%	5.60%	4.8%	5.1%	3.7%	3.6%
Poor self-rated health	5.5%	27.5%	5.70%	29.20%	6.7%	45.7%	21.6%	34.5%
Baseline years	2002, 2005		2002, 2005		2001, 2004, 2009		2001	

Notes: Full sample consists of working age adults in the labour force who are not in receipt of employment insurance, whereas analytic sample consists of only those individuals from the full sample who had reported variation in the health outcome over the study period. For the categorical variables, the associated numbers correspond to proportions. For continuous variables, numbers are means. Household income is in 2015 currency of each country, adjusted by household size. Ontario sample comprised those who remained in Ontario throughout the study period

^a Regions are as follows: for Canada: 1 = Atlantic, 2 = Prairies, 3 = Quebec, 4 = Ontario, 5 = British Columbia; for the US: 1 = Northeast, 2 = North Central/Midwest, 3 = South, 4 = West; for the UK: 1 = North East, 2 = North West, 3 = Yorkshire and Humber, 4 = East Midlands, 5 = West Midlands, 6 = East of England, 7 = London, 8 = South East, 9 = South West, 10 = Wales, 11 = Scotland, 12 = Northern Ireland

and “Have you yourself or jointly with others since [baseline year] received Jobseeker’s Allowance?” A “Yes” response to at least one of the questions was coded as exposed (1), and compared to those who received neither Income Support nor Job Seeker’s Allowance. The exposure periods across countries were thus roughly comparable.

Covariates By exploiting within-individual variation in exposure over time, we controlled for “fixed” (time-invariant) covariates, such as immigrant status, and birth cohort, as well as more difficult to measure variables such as anchoring of the response (i.e., differences in the consideration of one’s initial health in judging one’s future health, which can be particularly important

for subjective measures such as self-rated health) and intrinsic differences between individuals (e.g., unobservable differences between individuals that are associated with health status, such as genetic dispositions and personality traits).

We further added to our regression models observable, time-varying regressors. These included the following: labour force status, gender, age, education, visible minority status, home ownership, geographical area, marital/partnership status, household income, and number of children.

Statistical analyses We began by describing sample characteristics. Next, we explored the association between social assistance receipt and individual health status by exploiting the panel nature of our data using fixed effects modelling. The basic notion is that, by measuring whether or not social assistance was received in the baseline year, and assessing health during the baseline year, we can draw an association between receipt of social assistance and subsequent change in health status.

Our model specification is given by

$$\{y_{it} = \beta x_{it} + \gamma z_{it} + \nu_t + \alpha_i + \epsilon_{it}, \quad i = 1, 2, \dots, m \quad (1) \\ t = 1, 2, \dots, T$$

where i indexes individual respondents and t indexes time periods. y_{it} is a dichotomous indicator of poor health = 1 if an individual reported poor self-rated health and zero otherwise; x_{it} is a dummy variable = 1 if an individual's family received social assistance; z_{it} is a vector of time-varying control variables, including labour force status, family composition, and household income, and ν_t captures time-specific macro shock common to all respondents using a set of calendar-year dummies for the SLID and the BHPS samples, and survey-year dummies for the SIPP sample due to its independent short panels which do not overlap. Finally, α_i is an individual specific error component capturing time-invariant unobserved personal characteristics affecting individual health, which could also be correlated with social assistance receipt; ϵ_{it} is independently and identically logistic distributed individual specific error term. Consistent estimates of the coefficients are obtained by conditional maximum likelihood method. Stata's xtlogit, fe command is utilized for this estimation.

We first estimated a simple logistic regression on the pooled years of the samples as Model 1 controlling for all constant and time-varying individual characteristics. Model 2 fitted a crude fixed effects logistic model associating individual health outcome with an indicator variable for receipt of social assistance and unemployment status. We extended this baseline specification by adding controls for whether an individual lives in a couple-headed household and whether the individual has children in Model 3. We added a control for adjusted household income in Model 4.

The fixed effects logit estimator identifies the health effect by exploiting within-individual variations in receiving the treatment (social assistance) and health status over time. In fact, the coefficients are estimated on the sample of individuals who experienced change in health status at least once during the panel period. Individuals with unchanged outcome drop out of the conditional maximum likelihood function. Hence, the reported analytic sample sizes of the fixed effects logit models are substantially smaller than the full sample. There is therefore a trade-off between less bias of fixed effects logit specification and the resulting loss in efficiency.

We considered labour market status-specific differences in the effect of social assistance on poor self-rated health status. However, computing the marginal effects of the interaction term in the fixed effect logit model is impossible without making a further assumption that all the fixed effects are zero and the alternative, odds ratio interpretation, is discouraged due to its complications (Karaca-Mandic et al. 2012). Moreover, in the case of short panels, the maximum likelihood estimates of non-linear fixed effect models (such as logit fixed effects) are known to be biased and inconsistent due to the incidental parameter problem (Lancaster 2000). The former assumption can contradict the rationale for using the fixed effects approach. Hence for robustness check, we estimated a linear probability model with interaction terms controlling for individual fixed effects. To account for the sampling design of the surveys, we used survey weighting in all models. All individuals were weighted by their first-year time-constant weights.

Results

Ontario Table 1 presents a description of the baseline characteristics of the full and analytic samples. Relative to the full sample, those in our analytic sample from Ontario were more likely to be non-White (25.3% vs. 21.7%), more likely to be an immigrant (33.3% vs. 28.2%), and more likely to have completed less than a secondary level of education (16.1% vs. 10.7%). Relative to the full sample, they exhibited higher rates of poor self-rated health (27.5% vs. 5.5%), social assistance reciprocity (8.5% vs. 5.2%), and unemployment (6.2% vs. 4.0%). Those in the analytic sample also reported lower average household income compared to those in the full sample. Table S1 presents trends in the labour market status, social assistance coverage, and self-rated health of the full and analytic samples over the period of study. Within the analytic sample, there was a somewhat secular trend towards increasing rates of poor self-rated health, from 26.2% in 2002 to 40.8% in 2010. Rates of social assistance reciprocity differed from year to year, ranging from 7.2% to 11.0%.

Table 2 and Fig. 1 present the results of our pooled logit and fixed effects models for Ontario. In the pooled logit model

Table 2 Odds of reporting poor or fair self-rated health in Ontario: fixed effects modelling of survey of labour and income dynamics 2002–2010

	Pooled logit	Fixed effect		
	M1	M2	M3	M4
Received social assistance	2.380*** (1.798–3.150)	1.16 (0.691–1.947)	1.173 (0.699–1.968)	1.194 (0.713–2.000)
Unemployed	2.315*** (1.538–3.484)	2.241** (1.339–3.749)	2.219** (1.325–3.717)	2.373** (1.406–4.005)
Couple headed	0.819* (0.673–0.996)		0.814 (0.515–1.287)	0.758 (0.479–1.200)
Have children	0.826* (0.695–0.983)		0.818 (0.548–1.221)	0.833 (0.558–1.242)
Annual family income in thousand	0.986*** (0.983–0.990)			1.010* (1.002–1.017)
Year fixed effects	Yes	Yes	Yes	Yes
Number of obs	20,071	2701	2701	2701
Number of groups		608	608	608
Log likelihood	– 5441.03	– 1250.88	– 1249.76	– 1246.8
Chi-squared	425.06	107.24	109.49	115.4

Notes: Outcome modelled as poor or fair versus good, very good, or excellent. For Ontario, we look at those who remained in Ontario throughout the study period. In M1, we further controlled for age, gender, minority status, education, immigrant status, birth cohort, and home ownership. 95% confidence intervals are in parentheses

* $p < 0.05$; ** $p < 0.01$; *** $p < 0.001$

(Model 1), after controlling for a range of demographic and socio-economic variables, factors associated with poor self-rated health included being unemployed (OR 2.315, 95% CI 1.538–3.484) and receiving social assistance (OR 2.380, 95% CI 1.798–3.150), while being married or cohabitating, having children, and having a higher annual household income, were all associated with better self-rated health. In the crude fixed effects model (Model 2), moving into unemployment was

associated with poor self-rated health (OR 2.241, 95% CI 1.339–3.749). By contrast, the association between social assistance receipt and self-rated health was no longer significant. These results remained stable after controlling for time-varying characteristics (i.e., household structure and income) (Models 3–4).

Canada Table 1 presents a description of the baseline characteristics of the full and analytic samples. Relative to the full sample, those in our analytic sample from Canada were more likely to be non-White (23.4% vs. 17.9%), more likely to be an immigrant (25.6% vs. 19.9%), and less likely to have completed a post-secondary degree (17.4% vs. 22.7%). Relative to the full sample, they exhibited higher rates of poor self-rated health (27.5% vs. 5.5%), social assistance reciprocity (8.5% vs. 5.2%), and unemployment (6.2% vs. 4.0%). Those in the analytic sample also reported lower average household income compared to those in the full sample. Table 3 presents trends in the labour market status, social assistance coverage, and self-rated health of the full and analytic samples over the period of study. Within the analytic sample, there was a somewhat secular trend towards increasing rates of poor self-rated health, from 27.2% in 2002 to 42.5% in 2010. Rates of social assistance reciprocity differed from year to year, ranging from 5.6% to 8.6%.

Table 3 and Fig. 1 present the results of our pooled logit and fixed effects models for Canada. In the pooled logit

Fig. 1 Odds of reporting poor health, by social assistance status. Note: Fixed effects logit regression results controlling for time-invariant variables, unemployment status, household composition, and income; 95% confidence intervals are shown

Table 3 Odds of reporting poor or fair self-rated health in Canada: fixed effects modelling of survey of labour and income dynamics 2002–2010

	Pooled logit	Fixed effect		
	M1	M2	M3	M4
Received social assistance	2.023*** (1.705–2.401)	1.525* (1.102–2.110)	1.524* (1.101–2.109)	1.525* (1.101–2.110)
Unemployed	2.224*** (1.755–2.817)	1.614** (1.163–2.240)	1.619** (1.166–2.248)	1.620** (1.166–2.252)
Couple headed	0.830** (0.740–0.930)		0.799 (0.607–1.052)	0.798 (0.605–1.054)
Have children	0.893* (0.807–0.987)		1.008 (0.785–1.295)	1.008 (0.785–1.296)
Annual family income in thousand	0.988*** (0.986–0.991)			1.000 (0.995–1.005)
Year fixed effects	Yes	Yes	Yes	Yes
Number of obs	69,058	8870	8870	8870
Number of groups		1983	1983	1983
Log likelihood	– 14,551.22	– 3243.36	– 3242.05	– 3242.05
Chi-squared	1101.72	203.2	205.82	205.83

Notes: Outcome modelled as poor or fair versus good, very good, or excellent. In M1, we further controlled for age, gender, minority status, education, immigrant status, province, birth cohort, and home ownership. 95% confidence intervals are in parentheses

* $p < 0.05$; ** $p < 0.01$; *** $p < 0.001$

model (Model 1), after controlling for a range of demographic and socio-economic variables, factors associated with poor self-rated health included being unemployed (OR 2.224, 95% CI 1.755–2.817) and receiving social assistance (OR 2.023, 95% CI 1.705–2.401), while being married or cohabitating, having children, and having a higher annual household income, were all associated with better self-rated health. In the crude fixed effects model (Model 2), moving into unemployment (OR 1.614, 95% CI 1.163–2.240) and social assistance reciprocity (OR 1.525, 95% CI 1.102–2.110) were still associated with poor self-rated health. These results remained stable after controlling for time-varying characteristics (i.e., household structure and income) (Models 3–4).

US and UK Table 1 presents descriptive findings and supplementary tables (Table S2, Table S3) provide a discussion of descriptive analyses and also present detailed models for the US and the UK. Results of the final model (Model 4) are presented in Fig. 1 and suggest no significant effect of social assistance in the US, and higher odds of reporting fair or poor self-rated health when moving into social assistance in the UK (OR 1.294, 95% CI 1.005–1.666).

Supplementary analyses using linear probability modelling suggested no significant interaction between social assistance and employment status in any jurisdiction (see supplementary Tables S4–S7).

Discussion

The results of our analyses suggest that, contrary to what might be expected, social assistance programs in Canada and in peer countries are not succeeding at improving the health of the income-insecure. The health of social assistance recipients was worse (Canada, UK) or no different (Ontario, US) than the health of non-recipients. The use of fixed effects modelling provides confidence in the robustness of our inferences against alternative explanations. However, despite using a more rigorous method of estimation, our results are consistent with prior cross-sectional research on the association between social assistance and health, which also suggests social assistance is not succeeding in improving health (Jayakody et al. 2000).

On the surface, these results might suggest that providing benefits to the income-insecure is inherently problematic for health. But, from a plethora of theoretical and empirical findings, such an explanation is without merit (Basu et al. 2016; Marmot 2005; Irwin et al. 2006; Golberstein 2015). Rather, there are other far more plausible explanations. The amount of income provided by social assistance may be insufficient to buffer individuals and families in times of income scarcity and insecurity (Basu et al. 2016; Evidence Network n.d.; Curtis and Pennock 2006). Work conditionalities attached to receipt of social assistance may be offsetting health benefits of social assistance, because recipients are being exposed to precarious labour market conditions that are all too common in low-wage jobs (Segal 2012; Kim et al. 2008; Vives et al. 2013). The

stigma associated with receiving means-tested benefits may also work to counteract potential health effects of social assistance (Segal 2012; Iceland 2008; Meyers et al. n.d.).

Findings that Ontario and the US exhibited comparatively “better” results than Canada and the UK—with no difference between social assistance recipients and non-recipients, as opposed to a decline in health status among recipients—are especially curious. In particular, the US is characterized by even less generous benefits. We speculate that US results are a consequence of a selection process related to the more stringent eligibility criteria for receiving social assistance in the US (Farrants et al. 2016), which means many more socio-economically disadvantaged individuals are denied social assistance in the US compared to other liberal market economies, thus creating a larger pool of people with higher risk factors for ill health among the non-recipient group than in Canada or the UK. Put differently, “better” results in the US are less likely to be related to the more beneficial nature of social assistance there, and are more plausibly an artefactual consequence of the limited reach of US social assistance. This may also be true of Ontario, which may have differential selection into social assistance compared to the rest of Canada as a whole.

The results of our study suggest three main policy avenues to explore (Evidence Network n.d.). The first is to broaden eligibility criteria of social assistance, so that more people in need are covered. The second is to increase the generosity of benefits provided by social assistance. The third is to remove work conditionalities, which particularly in an era of precarious employment may expose social assistance recipients to additional stress and strain. The Ontario government recently announced social assistance reforms, which will change the structure of benefits and will impose more stringent work conditionalities. The health impact of this policy shift should be evaluated in the future.

By examining change in individuals over time, rather than comparing individuals, we controlled for many of the “fixed” characteristics, which differ between individuals and which can cause confounding. However, there are a range of time-varying characteristics for which, due to lack of data, we did not control, and which may have influenced our results. These include duration of unemployment or precarious employment, and unreported sources of income and wealth. These factors may influence selection into social assistance. We were also unable to examine a broader range of health outcomes. These are also avenues we recommend for future research.

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Compliance with ethical standards

Conflict of interest The authors declare that they have no conflict of interest.

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