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Abstract

The effect of insurance expansions on the distribution of health status is still a matter we know little about. This paper draws upon new measures of pure health inequality and mobility in health which accommodates categorical data to understand how an expansion of public insurance affects both inequality and mobility in health. These measures require a measure of health status that is either "upward-looking" or "downward looking". We find that, the distribution of health worsens in Mexico between 2002 and 2009, although the change is only consistent for the upward looking definition of status. Together with the lack of mobility in health observed, we can thus conclude that Mexico is becoming more rigid over time insofar as the distribution of health status.

JEL-Codes: I180.

Keywords: self-reported health, health inequality, health mobility, health insurance, Mexico.

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1. Introduction

The distribution of the overall attainment of health has become an important indicator to evaluate a country's health system performance (WHO 2000), as well as the success of policy interventions to extend insurance coverage. Nonetheless, measuring changes in the distribution of health, and specifically, changes in inequality and mobility in a population is far from straightforward. A growing number of studies have focused on both developing measurement tools and providing evidence for specific countries or groups of countries (van Doorslaer and Van Ourti 2011). However, most of these studies have only addressed health disparities across socioeconomic status through concentration indices. While this approach has been helpful in drawing attention to dimensions of well-being other than income, it raises some conceptual and methodological concerns that we attempt to address in this paper.

Approaches that focus on measuring socioeconomic inequalities in health are problematic on a number of grounds. First, it may well be argued that all health inequalities should be a cause of concern and not only those related to socioeconomic status (Gakidou et al. 2000). Second, the analysis of health-related inequalities often draws on unsatisfactory cardinalisation procedures to deal with ordinal variables such as self-assessed health (SAH). Finally, socio-economic measures do not address the fact that income and health might well be codetermined, as evidence suggests. So, with few exceptions (Contoyannis et al. 2004), studies that have focused on measuring health mobility also tend to focus on socioeconomic mobility. Alternative distributional measures of *pure* health inequality and mobility are less problematic and more suitable to evaluate the effect of policy interventions.

This paper employs a recently developed class of indices suitable for ordinal data to analyse the pattern of pure health inequality and mobility between 2002 and 2009 in Mexico (Cowell and Flachaire 2017, forthcoming). The Mexican case provides an especially suitable setting, as an ambitious health reform took place over that period through the implementation of the Seguro Popular (SP), a public health insurance programme created to provide access to a generous package of health services to those previously excluded from insurance. Indeed, by the beginning of the last decade health insurance coverage was segmented by labour status in Mexico. Only formal sector workers and their families had access to health care and other social security benefits; about half of the population went without insurance. According to administrative records, the health insurance extension managed to attain full universal health care coverage by 2012 (Knaul et al. 2012).

Insurance coverage, whether public or private, provides financial security, and specifically reduces the risk of unpredictable medical costs that households would otherwise absorb. If such costs are too high, individuals go without health care, which can have undesirable consequences for their health. Health insurance provides access to primary care and preventive services too. In particular, if coverage is provided to the entire population (as in the case of Seguro Popular), it could reduce pre-existing disparities in the access to health care inputs, and so reduce pure health inequalities. This would be expected to reduce the disparities in health across the population, and more specifically, to improve the health of those with the poorest health, hence improving pure health mobility.

Nonetheless, the production of health depends on a large list of inputs in addition to health care access. Moreover, the universalization of health insurance alone guarantees neither use of, nor access to needed health care, especially preventative services. Whether increased access takes place, in particular to high-value health care that improves health status, is an empirical question. Overall, the consensus from recent studies drawing on insurance extensions in the US is that coverage improves individuals' perceived health (see a summary in Sommers et al. 2017). This is exemplified by the Oregon study, a key and paradigmatic randomised expansion of health insurance in the US, that found a 25 per cent increase in the likelihood of individuals reporting good or very good health after one year (Finkelstein et al. 2012). The evidence on the effects of the Seguro Popular is more limited, but Teruel et al. (2012) also found that the programme increased the probability of reporting good health by 6 per cent. However, little is known about the effects on the distribution of health. Evidence for China, a country that has also undergone important reforms to increase insurance coverage, suggests that health insurance is associated with reductions in health inequalities, but the overall trend seems to be largely driven by factors outside the health system (Wang and Yu 2016). In fact, health inequalities have increased in China between 1997 and 2009 in both rural and urban areas. We expect to provide new evidence on the potential association between health insurance expansions and the distribution of health.

To fully exploit the information on individual changes in health status between points in time, we also analyse short-run mobility in health. According to Shorrocks' (1978) seminal paper in the income dynamics literature, the concept of mobility captures the extent to which inequality fades over time. Hence, the existence of health mobility would suggest that inequality declines could be expected in the long term. Likewise, a strong persistence in health would suggest that inequality declines are less likely. Again, we use a recently developed mobility indicator that allows dealing with ordinal variables (Cowell and Flachaire, forthcoming). A major contribution of this class of indices is that it separates the definition of status (i.e., the position in the distribution of health) from the definition of mobility.

The study is organised as follows. Section 2 contains a critical guide to the relevant literature. Section 3 describes the setting for the analyses, i.e., the expansion of public health insurance that took place in Mexico over the last decade. Section 4 describes the data, measures, and analytic sample. Sections 5 and 6 presents the methods employed to analyse both inequality and mobility in health and the results, including some robustness checks. Section 7 discusses the results.

2. Background

2.1. Health inequalities

The study of health inequalities has been the focus of numerous studies over the past decades. Most analytic tools employed in these studies have been inspired by the income inequality literature. But there are salient differences between the nature of income –an unbounded, cardinal variable— and health –commonly measured with a categorical variable, for which the real distance between the categories is unknown. In particular, concentration indices of health on income (CI) are the most popular tool to measure income related health inequalities (see a survey in Wagstaff and van Doorslaer 2000, and van Doorslaer and Van Ourti 2011). The World Bank has even published a practical guide to facilitate the estimation of CI (O'Donnell et al. 2008). One of the features that makes this measure attractive is that it can be decomposed into the contributions of a set of characteristics, provided the relevant outcome can be written as a linear function of these characteristics (Wagstaff et al. 2003). But as CI should only be used with cardinal variables, arbitrary cardinalisation methods have been commonly applied. For example, van Doorslaer and Jones (2003) use an ordered probit model to convert SAH categories into a continuous index that is then employed to measure inequality. According to Erreygers and van Ourti (2010), however, this rescaling procedure does affect the estimates.[1]

Another aspect that makes the CI approach problematic is that the analysis is based on a measure of status that ranks individuals according to socioeconomic status, i.e., individual status is given by their position in the income (or consumption) distribution, as opposed to a natural health ranking akin to pure health inequalities. The use of CI implies that all socioeconomic inequalities of any kind are considered illegitimate (unfair and avoidable), and so ignores the fact that some income differences across individuals may be a matter of choice itself or may reflect variations in preferences (Fleurbaey and Schokkaert 2011), and that income and health may be co-determined. Furthermore, the CI approach neglects other aspects of inequalities in health. While health disparities due to demographics such as age and sex are normally considered legitimate (hence the demographic standardisation of health status is a common practice), the role of other factors as a source of (legitimate/illegitimate) inequalities is ignored. Systematic health disparities have been found with respect to race, ethnic origin, place of residence, and other characteristics, however (e.g. King M et al. 2009, Cook et al. 2010). Therefore, it has been argued that all health inequalities should be a cause of concern and not only those related to socioeconomic status (Gakidou et al. 2000).

[1] Indeed, Costa-Font and Hernández (2013) show in a meta-regression analysis that most of the variation in health inequality estimates comes from differences in the cardinalisation of health status.

In this study, we use an approach to measure pure health inequalities before and after the Mexican health insurance expansion that overcomes the technical and conceptual difficulties outlined above. In particular, we estimate a class of indices that do not require any cardinalisation and use a similar status concept to those used in poverty and relative deprivation analyses (Cowell and Flachaire 2017; see section 5.1).

While the analysis of income inequalities has evidenced that Mexico is one of the most unequal countries (Esquivel 2015), little is known about the distribution of health. A few studies that have addressed this issue, have employed the most common CI approach and have mainly focused on health care (Urquieta-Salomón and Villarreal 2016, Barraza-Lloréns et al. 2013). In the case of China, a country that has also recently increased health insurance coverage, the study of the distribution of health has received much more attention (e.g. Baeten et al. 2013, Tang et al. 2008), but again most analyses have focused on income-related health inequalities. The study by Wang and Yu (2016) is an exception that finds that health inequality considerably increased between 1997 and 2009 in China. The authors argue that this is likely related to factors outside the health system, such as increasing income inequality and poverty, and environment deterioration. In fact, their results suggest that health insurance contributed to the reduction of health inequalities, although the overall pattern was in the opposite direction. This study will help to shed light on this finding using data on Mexico.

2.2. Health dynamics

Health dynamics have been much less studied than health inequalities. Hauck and Rice (2004) and Contoyannis et al. (2004) are relatively recent exceptions that rely on measurement tools employed in the income dynamics literature. Hauck and Rice (2004) use variance components random effects models and linear dynamic regression models to analyse mobility in a cardinal indicator of mental health taken from the British Household Panel Survey (BHPS). In the first case, the measure of mobility is obtained from the proportion of the total variability in health attributed to the permanent component (i.e., unobserved individual heterogeneity); in the second case, the estimated coefficient of the lagged health variable indicates the extent of mobility. They find there is much mobility in mental health but this varies across socioeconomic groups. In particular, the incidence and persistence of mental illness is higher among low income individuals. Contoyannis et al. (2004) also use a dynamic regression approach with data from the BHPS. Since their health measure is a categorical indicator of SAH, however, their specification is non-linear (namely, a dynamic panel ordered probit). Unlike Hauck and Rice (2004), they provide evidence of substantial health persistence and hence limited pure health mobility. Additionally, they show that attrition does not alter their findings.

While these studies are important to assess the existence of mobility in health, a different approach is needed if the objective is to analyse mobility patterns. Here we use a class of measures to compare mobility during the first half of the Mexican health insurance expansion with mobility during the second half of the expansion (see section 5.2).

3. The expansion of health insurance in Mexico

Before the most recent reform, health insurance coverage in Mexico was attached to formal employment. As a consequence, informal sector workers and their families, who account for approximately half of the population, were uninsured. This is a major problem in low- and middle-income countries, which is typically addressed by extending or subsidising insurance coverage.

Formal sector workers and their families have access to health services provided by social security institutions.[2] These institutions have their own facilities and are centrally managed by the federal government. Their funding comes from payroll taxes, employer contributions, and general revenues; no co-payments apply. The uninsured, on the other hand, had access to public facilities that are funded through general revenues and administered by state governments, but a scheme of fees based on self-reported income applied.

[2] There are several private providers of health services in Mexico, but since only 3 per cent of the population have private insurance (OECD 2005), these are mostly funded through out-of-pocket expenditure.

Although the government is the provider of health services through both social security institutions and Ministry of Health facilities, public resources were historically skewed towards the former. In 2000, public per capita expenditure for the uninsured was MX$1,482.4, while the corresponding figure for social security beneficiaries was more than double (MX$3,197.5) (Secretaría de Salud 2013).[3] This resulted in large differences in quality and large out-of-pocket expenditures.

The SP was created to guarantee access to health care as a universal right. Accordingly, the only eligibility criterion for the SP is not being a beneficiary of social security. Also, the benefit package guarantees access to a wide range of preventive and treatment interventions that cover most causes of morbidity and mortality (González-Pier et al. 2006). Moreover, several services have been added over the years; between 2004 and 2012, the interventions offered increased from 91 to 284. The government estimates that these interventions cover 100 per cent of the demand of primary care and 85 per cent of the demand for hospitalisation and surgery (CNPSS 2015).

The allocation of public resources also radically changed with the implementation of the SP. According to the rules of the programme, it should be financed through federal contributions (composed by the *cuota social*, which is an annual transfer equivalent to 3.9 per cent the minimum wage per beneficiary, plus an additional transfer of 1.5 times the *cuota social*), state contributions (0.5 times the *cuota social*), and progressive contributions from beneficiaries. In practice, however, the contributions from beneficiaries are negligible (less than 1 per cent the SP annual budget; CNPSS 2015), so the programme is essentially financed with general revenues. By 2011, the gap in public per capita expenditure between those with and without social security beneficiaries had narrowed. Also, public health expenditure grew from 2.6 per cent of GDP in 2000 to 3.1 per cent in 2011 (Secretaría de Salud 2013).

[3] Figures in constant pesos. Health expenditure data are publicly available on the Federal and State Health Accounts System (Sicuentas) administered by the Ministry of Health.

The SP started as a pilot in 2002 with 1.1 million beneficiaries distributed across 341 municipalities in 20 states.[4] By 2007, all the municipalities had at least one affiliate, and the total coverage had increased to 21.8 million individuals. In 2012, it was formally announced that the country had reached universal coverage (Knaul et al. 2012); the programme records indicate that the coverage reached 52.9 million beneficiaries in that year.

The studies that have analysed various aspects of the SP typically indicate that the programme has reduced health expenditures; there is also some evidence on the positive effects on utilisation and health (see a comprehensive review in Knaul et al. 2012). Teruel et al. (2012) show that the programme increased self-assessed health by 6 per cent. Distributional aspects have received little attention, however, except for the assessment of financial impacts (e.g. King G et al. 2009). In general, the analysis of health disparities is scarce in Mexico despite being an important topic in the political agenda. To examine whether SP managed to change the distribution of health in the Mexican population is the main purpose of this study.

4. Data

4.1. The Mexican Family Life Survey

The Mexican family Life Survey (MxFLS) is a longitudinal survey covering the past decade. Three waves are available.[5] The first was collected in 2002, before the formal onset of health insurance expansion in Mexico; the second was collected between 2005 and 2006, when coverage levels of the Seguro Popular were between 11 per cent and 15 per cent; and the third wave was collected between 2009 and 2010, when the programme's coverage had reached nearly 40 per cent of the population.[6]

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[4] The 32 states that comprise the country are divided into municipalities, which are the smallest autonomous political entities. There are currently 2,457 municipalities.

[5] All the data bases, questionnaires, and supplementary information of the MxFLS are available in Spanish and English at http://www.ennvih-mxfls.org. Rubalcava and Teruel (2006, 2008, 2013), also available at the MxFLS's website, describe the planning and design of the survey, as well as the content and structure of the data sets.

[6] A few households were interviewed for the second and third round in 2007 and 2011-13, respectively. Coverage levels are based on information from administrative records of the Seguro Popular programme and population figures from the National Institute of Geography and Statistics (INEGI).

The MxFLS employed probabilistic, stratified, and multi-staged sampling design, and is representative at the national level, for rural and urban areas (less than 2,500 inhabitants and 2,500 inhabitants or more, respectively), and for five regions: south-south east, centre-occident, centre, northeast, and northwest.[7] The first wave included approximately 8,440 households and more that 35,000 individuals distributed among 150 communities throughout Mexico.

The information collected in the MxFLS covers a wide variety of topics. Indicators of expenditure, land use, economic shocks, and violence and victimisation, among others, are provided at the household level. Other information such as education, labour supply, marital and fertility history, migration history, time allocation, health status, health care utilisation, and cognitive ability is collected at the individual level. Finally, qualitative and quantitative information at the community level is also available, including commercial infrastructure and education, health and transportation services, and prices of goods and services.

The MxFLS interviews were implemented as follows. One or two adults reported all the information related to the socioeconomic status and demographic composition of the household. In parallel, each household member 12 years and older was interviewed to collect the information at the individual level. The information for children under 12 years was provided by an adult member of the household (their primary caregiver if possible). If any adult 15 years and older was not present at the moment of the interviews, proxy information was collected from other household members. This information is reported in a separate book so it can be easily identified.

[7] These regions correspond to the regions defined in the National Development Plan (*Plan Nacional de Desarrollo*) for 2000-2006 and are defined as follows: 1) the south-south east region covers the states of Campeche, Yucatán, Chiapas, Oaxaca, Quintana Roo, Tabasco, Guerrero, and Veracruz; 2) the centre-occident region covers the states of Jalisco, Michoacán, Colima, Aguascalientes, Nayarit, Zacatecas, San Luis Potosí, Guanajuato; 3) the centre region covers the states of Mexico City, Querétaro, Hidalgo, Tlaxcala, Puebla, Morelos, and Mexico; 4) the northeast region covers the states of Tamaulipas, Nuevo León, Coahuila, Chihuahua, and Durango; and 5) the northwest region covers the states of Baja California, Baja California Sur, Sonora, and Sinaloa.

4.2. Measures

The health variable employed is the response to the question *currently, do you consider your health is…?*, originally coded as very good (1), good (2), regular (3), bad (4), and very bad (5). This information is available for individuals 15 years and older, which constitute the basic unit of analysis. To calculate inequality indicators, the variable was recoded so that higher values represent better health (i.e., very bad health was recoded as 1, bad health was recoded as 2, and so on).

SAH information has been widely used in the literature that analyses the relationship between health and socioeconomic status (e.g. Adams et al. 2003, Deaton and Paxson 1998, Salas 2002), as well as in the studies that focus on the relationship between health and lifestyles (e.g. Contoyannis and Jones 2004). While SAH is a simple subjective indicator that provides an ordinal ranking of perceived health status, previous studies have shown that it is a good predictor of subsequent use of medical care (e.g. van Doorslaer et al. 2002) and subsequent mortality (e.g. Burström and Fredlund 2001). However, some studies have suggested that SAH may be measured with error if different groups of the population systematically consider different cut point levels when reporting SAH (Groot 2000, Sadana et al. 2000, Murray et al. 2001). Using SAH information from the Canadian National Population Health Survey, Lindeboom and van Doorslaer (2003) found that cut points varied with sex and age, although not with income and education. Our analysis of inequalities is therefore conducted for different population groups defined by sex, age, and type of residence area.

Proxy information of SAH was used if available to increase the sample size (see a detailed discussion of the sample size and the effect of non-response in section 4.3). Since this could be a potential source of bias due to the subjective nature of the variable, section 6.3 discusses the implications.

Other variables employed in the analyses include sociodemographic characteristics at baseline, namely, binary variables to indicate whether the individual was female, lived in rural areas, and was active in the labour market in the past 12 months. The region of residence, age group (15 to 30 years, 31 to 45 years, and 46 or older), education level as defined by the highest level of education completed (none, primary, secondary, high school, and university), marital status (cohabitating couple, separated or divorced, single, and widowed), and household size are also used.

4.3. Sample description

Like other longitudinal surveys, the MxFLS suffers from different types of non-response. *Attrition*, a type of non-response specific to longitudinal surveys, occurs when baseline participants are not able or willing to participate in subsequent waves of the survey. The reasons behind attrition can be death, serious illness, national or international migration, or simple refusal (see Uhrig 2008 for a review of the reasons of panel attrition). *Item non-response*, on the other hand, occurs when participants have missing information in some parts of the survey. This type of non-response may be caused by unwillingness to provide information that is considered sensitive, or simply because the answer is unknown. The survey context is also important to understand item non-response (Frick and Grabka 2005); the complexity of surveys like the MxFLS may well explain at least part of this problem. In either case, if non-response is completely random, the results would be unaffected and simple case-wise deletion would be a valid alternative (Rubin 1987). This is unlikely, however, and therefore constitutes a potential source of bias that must be addressed.

In practice, weighting and imputation methods are the most common ways of dealing with attrition and item non-response, respectively (Jenkins 2011). While specific weights can be constructed to address the research question of interest (e.g. Jenkins 2009, Contoyannis et al. 2004, Jones et al. 2006), those normally provided by the survey administrators are often employed. These weights are designed to produce estimates that represent the population from which the sample was drawn and to adjust for non-response. Imputed data is also frequently provided by the survey administrators, especially for variables such as income, with relatively high item non-response rates.

Both weighting and imputation, however, normally rely on different assumptions about non-response patterns that have to be considered. Many studies on income inequality and mobility have found that differential attrition does not have a substantial impact on the conclusions (see Jenkins 2011). Using the BHPS and the

European Community Household Panel to analyse socioeconomic determinants of health, Jones et al. (2006) also show that health-related attrition has little impact on the results. Frick and Brabka (2005), on the other hand, show that using only non-imputed data from the German Socioeconomic Panel Study significantly underestimates income mobility.

Overall, the MxFLS has relatively low levels of attrition. In particular, 9.2 per cent of the participants 15 years and older at baseline was lost to follow-up at wave 2, while an additional 7.3 per cent was lost to follow-up at wave 3 (table 1). Similar longitudinal surveys for other low- and middle-income countries have attrition rates above 16 per cent (e.g. the Vietnam Living Standards Measurement Survey and the Côte d'Ivoire Living Standards Measurement Survey; Falaris 2003). An exception is the Indonesian Family Life Survey, which served as a model for the MxFLS and has a comparable attrition rates (Thomas 2012). General, response rates of the MxFLS are also good but vary across books. For example, the book that contains information about household consumption has a response rate of 95 per cent at baseline, while the book on adult cognitive ability has a response rate of 85 per cent. Non-response in SAH, however, is relatively high. If no proxy responses are considered, between 17 per cent and 22 per cent of the participants have missing SAH information; only after considering proxy responses the item non-response decreases to 10 per cent approximately (table 1). In contrast, SAH non-response in the BHPS is less than 1 per cent (Lynn 2006).

If we consider both participation in all three waves and complete SAH information (including proxy responses), we end up with a balanced sample of 15,088 individuals or 45,264 wave-individual observations, which constitutes the main analytic sample. The weights of the MxFLS provided are used in the main analysis, as these adjust for non-response.[8] Section 6.3, however, explores other specifications to

[8] The survey materials available at the MxFLS website include a document that explains the calculation of the weights. In sum, the weights are first calculated at the household level as the inverse of the joint probability of selecting this last sampling unit. These weights imply three types of adjustments to account for non-response, for projections to the entire population, and for calibration. Once the

assess the robustness of the results, including unweighted estimates, multiple imputation, non-proxy information only, and inequality estimates using the unbalanced sample.

Table 1. MxFLS non-response. Sample of individuals 15 years and older at baseline

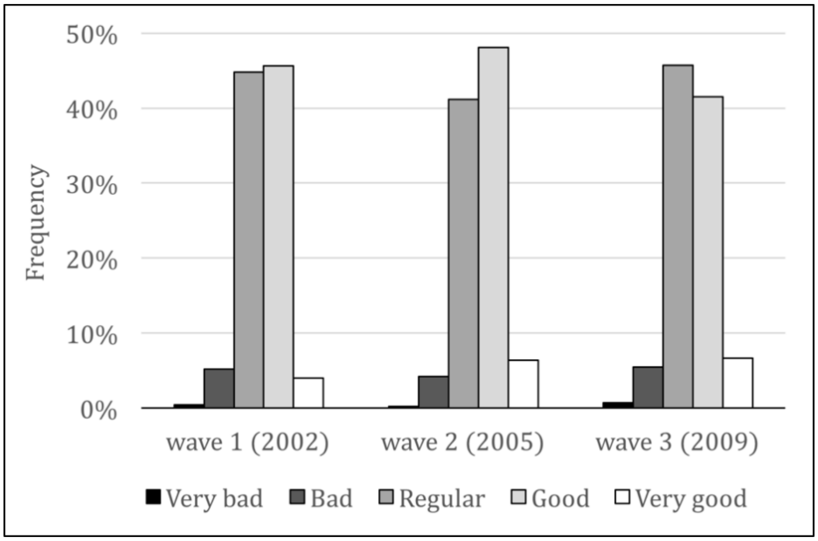
Wave	Individuals	Survival rate	Drop-outs	Attrition rate	Complete SAH information		Item non-response
					(no proxy)	(with proxy)	
1	23,724				19,778	21,610	8.9%
2	21,550	90.8%	2,174	9.2%	16,936	19,091	11.4%
3	19,971	84.2%	1,579	7.3%	15,546	17,635	11.7%

Notes: Only baseline participants considered (individuals added to replenish the sample in waves 2 and 3 are excluded from the analyses). 53 observations with no information of age at baseline are also excluded.
Source: Own estimates based on the Mexican Family Life Survey (MxFLS).

Figure 1 shows the distribution of SAH by survey wave. Most of the individuals have regular or good health. The share of those reporting very bad and bad health is slightly higher in the last wave, but the difference is not clear enough to claim that the distribution became worse over the period. Table 2 shows the sociodemographic characteristics of the sample. About half of these individuals were female, active in the labour market, and lived in rural areas at baseline. Their education level was generally low (11 per cent reported no formal education and 43 per cent had only completed primary education), and nearly two thirds lived with their couple. The sample is roughly equally distributed across the five regions.

household weights are adjusted, individual weights for each book (with and without proxy responses) are calculated.

Figure 1. Self-assessed health by survey wave



Notes: Respondents who participated in all three waves and have complete SAH (proxy information considered), n=15,088; unweighted percentages.
Source: Own estimates based on the Mexican Family Life Survey (MxFLS).

Table 2. Baseline characteristics of the balanced sample (unweighted figures)

	%/mean	n
Age (mean)	37.8	15,088
15 to 30 years	39.2%	5,913
31 to 45 years	30.5%	4,597
46 years or more	30.3%	4,578
Female	55.0%	8,298
Marital status		
Cohabitating couple	63.5%	9,584
Divorced or separated	3.9%	588
Single	28.1%	4,244
Widowed	4.4%	670
Highest education level completed		
None	10.9%	1,633
Primary	43.0%	6,472
Secondary	26.1%	3,924
High School	13.1%	1,972
University	7.0%	1,048
Worked in the past 12 months	53.5%	8,078
Household size (mean)	5.0	15,088
Rural	47.4%	7,154
Region		
South-south east	21.4%	3,228
Centre-occident	19.7%	2,965
Centre	18.3%	2,758

Northeast	19.5%	2,943
Northwest	21.2%	3,194

Notes: The balanced sample include respondents who participated in all three waves and have complete SAH (proxy information considered), n=15,088.
Source: Own estimates based on the Mexican Family Life Survey (MxFLS).

5. Methods

5.1. Measuring inequality in health

To analyse inequality in health status, we employ the Cowell and Flachaire (2017) inequality measure specifically developed to deal with ordinal variables such as SAH. Let n_k be the number of persons in each SAH category k= 1, 2,..., 5, where 1 is the least desired category (very bad health) and 5 is the most desired category (very good health). Then, the status of individual i who is in category $k(i)$ must be a function of either:

$$\sum_{l=1}^{k(i)} n_l \qquad \text{or} \qquad \sum_{l=1k(i)}^{K} n_l \tag{1}$$

Normalising by the size of total population, $n = \sum_1^K n_k$, so that individual's status is between 0 and 1 we have:

$$s_i = \frac{1}{n}\sum_{l=1}^{k(i)} n_l \qquad \text{or} \qquad s_i' = \frac{1}{n}\sum_{l=1k(i)}^{K} n_l \tag{2}$$

where s_i and s_i' are the downward and upward looking definitions of individual's status, respectively. If there was perfect equality, all the individuals would be in the same category and both expressions would be equal to one; this maximum-status is the reference point.

Based on a set of elementary axioms, Cowell and Flachaire (2017) show that inequality must take the form of an index in the following class:

$$I_\alpha(s) = \frac{1}{\alpha[\alpha-1]}\left[\frac{1}{n}\sum_{i=1}^{n} s_i^\alpha - 1\right], \alpha \in \mathbb{R}, \alpha \neq 0,1 \tag{3}$$

where $\alpha < 1$ indicates the sensitivity of the index to different parts of the health distribution. In particular, high values of α produce indices that are more sensitive to high-status inequality, while low and negative values produce indices that are more sensitive to low status. Depending on whether we use the definitions of status s_i or s_i' (2) to calculate $I_\alpha(s)$ (3), we will have an index of ordinal inequality based on a downward or upward looking status concept. The limiting form for the case where $\alpha = 0$ is:

$$I_0(s) = -\frac{1}{n}\sum_{i=1}^{n} log s_i \tag{4}$$

As can be seen in equations 3 and 4, the Cowell and Flachaire class of indices is actually similar to the well-known Generalised Entropy class of inequality measures GE_α (Cowell 1980, Shorrocks 1980). The second, however, takes the mean $\mu(s)$ as the reference point, which makes sense only if the measure of status is cardinal. Therefore, when ordinal variables such as SAH are employed, a common approach is to use an arbitrary cardinalisation to estimate GE_α. We use this approach in section 6.3 to test whether these results differ from those obtained using the Cowell and Flachaire inequality measures.

$$GE_\alpha(s) = \frac{1}{\alpha[\alpha-1]}\left[\frac{1}{n}\sum_{i=1}^{n}\left[\frac{s_i}{\mu(s)}\right]^\alpha - 1\right], \alpha \in \mathbb{R}, \alpha \neq 0,1 \tag{5}$$

$$GE_0(s) = \frac{1}{n}\sum_{i=1}^{n} log\frac{s_i}{\mu(s)} \tag{6}$$

$$GE_1(s) = \frac{1}{n}\sum_{i=1}^{n}\frac{s_i}{\mu(s)} log\frac{s_i}{\mu(s)} \tag{7}$$

Percentile bootstrap with 1,000 replications is used to calculate confidence intervals, i.e., we generate 1,000 bootstrap samples by resampling with replacement from the observed data, and then we estimate I_α^b (or GE_α^b), with b=1,...,1000, for each bootstrap sample. The percentile confidence interval is then:

$$CI_{percentile} = \left[c^b_{0.025}, c^b_{0.975}\right] \tag{8}$$

where $c^b_{0.025}$ and $c^b_{0.975}$ are the 2.5th and 97.5th percentiles of the Empirical Distribution Function of the bootstrap statistics.

All the routines to estimate the Cowell and Flachaire indices were programmed in Stata 14.2. The routine created by Jenkins to estimate generalised entropy measures was also used (Jenkins 2008).

5.2. Measuring mobility in health

Transition matrices or contingency tables provide a simple alternative to explore mobility. These matrices have been widely used to analyse mobility with categorical data such as employment status, educational attainment, or income quintiles (e.g. Ferrie 2005, Corak and Piraino 2010). Let S denote the set of all possible health status values, with S=[0,1] and subsets $S_1,\dots,S_k \subset S$ such that $U_{k=1}^K S_k = S$. Also, let n_{kl} be the number of individuals in S_k at time t_0 and S_l at time t_1. The transition matrix P is therefore a $K \times K$ array with elements

$$p_{kl} = \frac{n_{kl}}{\sum_{j=1}^K n_{kj}} \tag{9}$$

If nobody remains in the same position (perfect mobility), all the elements in the diagonal are equal to zero; if everybody stays in the same position (no mobility), all the elements in the diagonal are equal to one.

Mobility indices, however, provide a more useful approach that takes advantage of all the available information at the individual level. In particular, we use the Cowell and Flachaire (forthcoming) mobility index that has at least two important advantages compared to other commonly used mobility measures, namely, it is able to capture nonlinear relationships, and it separates the definition of individual's status from the definition of mobility.

Let u_i and v_i denote the status of individual *i* at time t_0 and time t_1, respectively, where $u_i, v_i \in S$ and S=[0,1], then the profile $z \coloneqq \{(u_i, v_i)_{i=1,\ldots,n}\}$ contains all the information about mobility for the population of *n* individuals. Based on a set of axioms on mobility orderings over all possible pairs *z*, Cowell and Flachaire (forthcoming) derived the following class of mobility measures that are independent of the population size and the scale of status:

$$M_\alpha = \frac{1}{\alpha[\alpha-1]n}\sum_{i=1}^{n}\left[\left[\frac{u_i}{\mu_u}\right]^{\alpha}\left[\frac{v_i}{\mu_v}\right]^{1-\alpha} - 1\right], \alpha \in \mathbb{R}, \alpha \neq 0,1 \tag{10}$$

where μ_u and μ_v are the means of u ad v, respectively, and α is a sensitivity parameter that characterises the particular members of the class. Positive values of α produce indices that are sensitive to downward movements, while negative α's produce indices that are sensitive to upward movements. The limiting forms for the cases where $\alpha = 0$ and $\alpha = 1$ are, respectively:

$$M_0 = -\frac{1}{n}\sum_{i=1}^{n}\frac{v_i}{\mu_v} ln\left(\frac{u_i}{\mu_u}/\frac{v_i}{\mu_v}\right) \tag{11}$$

$$M_1 = \frac{1}{n}\sum_{i=1}^{n}\frac{u_i}{\mu_u} ln\left(\frac{u_i}{\mu_u}/\frac{v_i}{\mu_v}\right) \tag{12}$$

Since we are employing an ordinal measure of health, proportions are used to define status: [9]

$$u_i = \widehat{F_0}(x_{0i}), \text{ and } v_i = \widehat{F}_1(x_{1i}) \tag{13}$$

[9] In other contexts (e.g. analyses of income mobility), different status concepts may be derived from a given data and the class of mobility measures M_α can be calculated for each status concept. Therefore, equations 3 to 5 can be actually considered a "superclass" of mobility measures (Cowell and Flachaire, forthcoming).

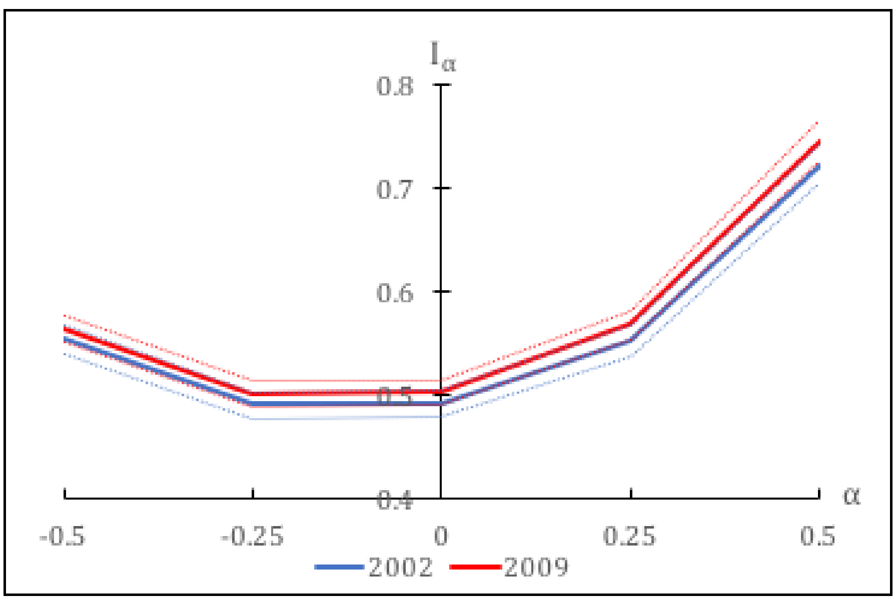
where $\widehat{F_K}(x) = \frac{1}{n}\sum_{j=1}^{n} I(x_{kj} \leq x)$ is the empirical distribution function of individual health in periods k=1, 2, and I(.) is an indicator function equal to 1 if its argument is true and equal to 0 otherwise. Percentile bootstrap with 1,000 replications is also used to calculate confidence intervals.

6. Results

6.1. Inequality in health

Figures 2 and 3 show the results of the Cowell and Flachaire inequality index (equations 3 and 4), using both the downward and upward looking definitions of status. In both cases, the point estimates suggest that health inequality increased between 2002 and 2009, but this change is only statistically significant when the upward looking definition is used. If we hold constant the definition of status, the conclusion is the same for different values of the sensitivity parameter α. Therefore, only the adoption of different status definitions affects the conclusions.

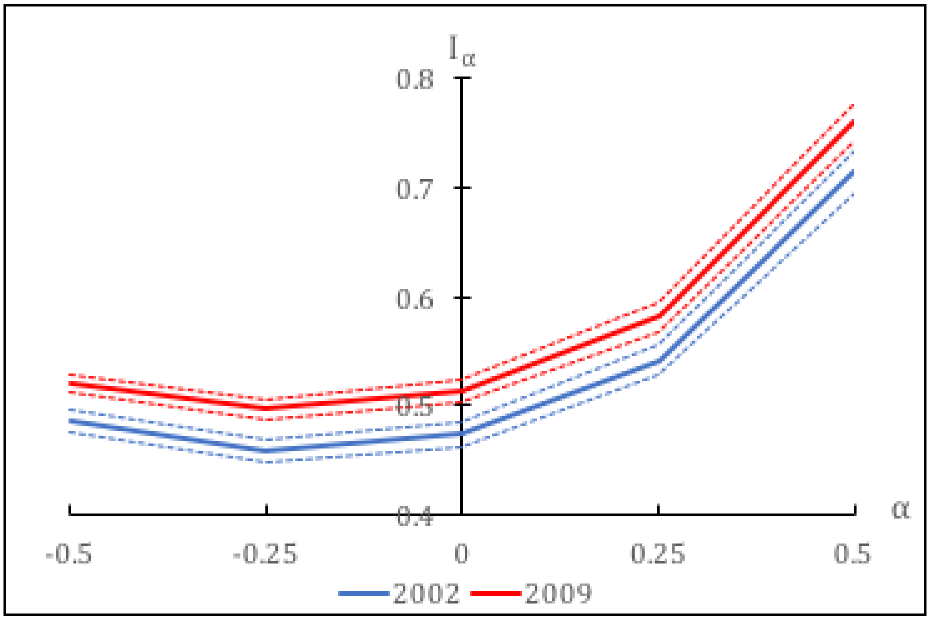
Figure 2. Health inequality during the public insurance expansion in Mexico. Downward looking status (balanced sample, weighted estimates)



Notes: Inequality is measured with the Cowell and Flachaire inequality index; the dotted lines represent 95 per cent confidence intervals estimated using bootstrap with 1000 replications. The

balanced sample includes the respondents who participated in all three waves of the MxFLS and have complete SAH information (proxy information considered); n=15,088.
Source: Own estimates based on the Mexican Family Life Survey (MxFLS).

Figure 3. Health inequality during the public insurance expansion in Mexico. Upward looking status (balanced sample, weighted estimates)



Notes: Inequality is measured with the Cowell and Flachaire inequality index; the dotted lines represent 95 per cent confidence intervals estimated using bootstrap with 1000 replications. The balanced sample includes the respondents who participated in all three waves of the MxFLS and have complete SAH information (proxy information considered); n=15,088. Source: Own estimates based on the Mexican Family Life Survey (MxFLS).

Table 3 further analyses health inequality among population subgroups holding the sensitivity parameter at $\alpha = 0$. As explained above (section 4.2), this is to account for the possibility that different groups of the population consider different cut point levels when reporting SAH, but also to assess whether inequality patterns vary among these population groups. The results obtained are similar to the results for the total population. If the downward looking definition of status is employed, inequality in health seems stable across both rural and urban areas, males and females, and cohorts, but increasing if the upward looking version is considered. The only group for which the increase is not statistically significant even if the upward looking status definition is used is the older cohort.

Table 3. Health inequality during the public insurance expansion in Mexico by baseline characteristics. Balanced sample, weighted estimates (sensitivity parameter $\alpha = 0$)

	Downward looking status			Upward looking status			
	2002	2005	2009	2002	2005	2009	n
Total	0.492	0.503	0.504	0.474	0.497	0.513	15,088
	[0.479, 0.505]	[0.490, 0.515]	[0.491, 0.516]	[0.461, 0.485]	[0.484, 0.508]	[0.502, 0.524]	
Area of residence							
Urban	0.485	0.504	0.507	0.471	0.499	0.514	7,934
	[0.469, 0.500]	[0.487, 0.519]	[0.492, 0.523]	[0.455, 0.486]	[0.484, 0.513]	[0.501, 0.527]	
Rural	0.497	0.499	0.492	0.473	0.485	0.505	7,154
	[0.481, 0.512]	[0.483, 0.514]	[0.476, 0.509]	[0.458, 0.487]	[0.470, 0.499]	[0.491, 0.518]	
Sex							
Male	0.489	0.500	0.510	0.466	0.485	0.511	6,790
	[0.469, 0.506]	[0.482, 0.517]	[0.490, 0.527]	[0.446, 0.483]	[0.467, 0.501]	[0.494, 0.527]	
Female	0.490	0.501	0.497	0.477	0.502	0.511	8,298
	[0.472, 0.508]	[0.483, 0.518]	[0.480, 0.516]	[0.460, 0.493]	[0.486, 0.516]	[0.494, 0.525]	
Age							
15-30 years	0.452	0.481	0.487	0.439	0.462	0.485	5,913
	[0.430, 0.470]	[0.461, 0.502]	[0.467, 0.507]	[0.417, 0.460]	[0.442, 0.481]	[0.464, 0.503]	
31-45 years	0.474	0.483	0.470	0.465	0.497	0.508	4,597
	[0.448, 0.496]	[0.458, 0.505]	[0.442, 0.494]	[0.442, 0.486]	[0.474, 0.516]	[0.490, 0.527]	
46+ years	0.500	0.491	0.496	0.493	0.496	0.503	4,578
	[0.475, 0.524]	[0.465, 0.514]	[0.470, 0.521]	[0.472, 0.512]	[0.474, 0.514]	[0.482, 0.521]	

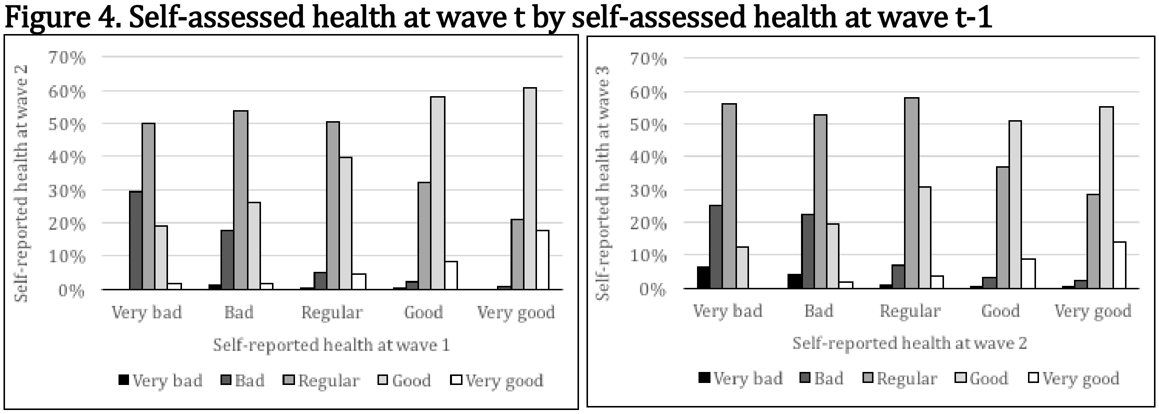
Notes: Inequality is measured with the Cowell and Flachaire index; 95 per cent confidence intervals, estimated using bootstrap with 1000 replications, are in brackets. The balanced sample includes the respondents who participated in all three waves of the MxFLS and have complete SAH information (proxy information considered).
Source: Own estimates based on the Mexican Family Life Survey (MxFLS).

6.2. Mobility in health

The results presented so far indicate that the distribution of health remained stable in Mexico during the past decade, or probably worsen (became more unequal) according to one of the definitions of status employed. Now we exploit individual changes in health between points of time to analyse mobility. In particular, we are interested in the extent to which health status in the previous period affects the distribution of health in the current period.

Figure 4 shows the distribution of SAH at wave 2 (or 3) by SAH at wave 1 (or 2). It seems clear that it is more likely to stay in the same state than to transition to another, especially if we look at the extreme categories. Those with very good health at wave 1, for example, are more likely to have very good health at wave 2. Similarly, those with very bad health at wave 2 are more likely to have very bad health at wave 3. The transition matrices in table 4 present an alternative way of analysing this. The rows indicate health in the previous period, while the columns indicate health in the current period. In general, the larger percentages are located in the diagonal or close to the diagonal, which is also an indicator of persistence in health. Additionally, we can see that the values in the diagonal that correspond to lower categories of health increased, but those that correspond to upper categories decreased. This suggests that overall mobility was likely stable.

Figure 4. Self-assessed health at wave t by self-assessed health at wave t-1



Notes: Unweighted percentages using the balanced panel (respondents who participated in all three waves of the MxFLS and have complete SAH information); n=15,088.
Source: Own estimates based on the Mexican Family Life Survey (MxFLS).

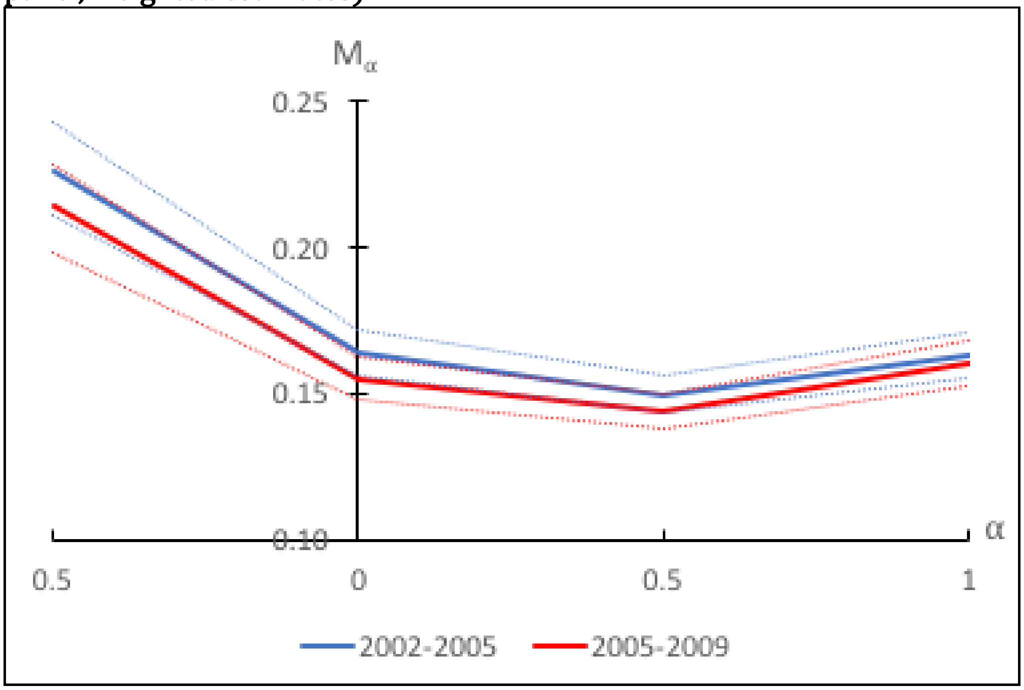
Table 4. Transition matrices, self-assessed health in Mexico

		2005					
		Very bad	Bad	Regular	Good	Very good	n
2002	Very bad	0.0%	29.3%	50.0%	19.0%	1.7%	58
		-	(19.0, 42.3)	(37.3, 62.7)	(10.8, 31.2)	(0.2, 11.4)	
	Bad	1.0%	17.5%	54.0%	25.9%	1.7%	784
		(0.5, 2.0)	(15.0, 20.3)	(50.4, 57.4)	(22.9, 29.1)	(1.0, 2.8)	
	Regular	0.3%	5.0%	50.7%	39.6%	4.3%	6,755
		(0.2, 0.5)	(4.5, 5.6)	(49.5, 51.9)	(38.5, 40.8)	(3.9, 4.8)	
	Good	0.1%	1.9%	32.0%	58.0%	8.0%	6,887
		(0.0, 0.2)	(1.6, 2.3)	(30.9, 33.1)	(56.8, 59.2)	(7.4, 8.7)	
	Very good	0.0%	0.8%	21.0%	60.6%	17.5%	604
		-	(0.3, 2.0)	(18.0, 24.5)	(56.6, 64.4)	(14.7, 20.8)	
	n	32	631	6,210	7,251	964	15,088
		2009					
		Very bad	Bad	Regular	Good	Very good	n
2005	Very bad	6.3%	25.0%	56.3%	12.5%	0.0%	32
		(1.5, 22.2)	(12.9, 42.9)	(38.7, 72.3)	(4.7, 29.3)	-	
	Bad	3.8%	22.5%	52.6%	19.3%	1.7%	631
		(2.6, 5.6)	(19.4, 25.9)	(48.7, 56.5)	(16.4, 22.6)	(1.0, 3.1)	
	Regular	0.8%	6.9%	58.1%	30.7%	3.5%	6,210
		(0.6, 1.0)	(6.3, 7.6)	(56.9, 59.3)	(29.6, 31.9)	(3.1, 4.0)	
	Good	0.4%	3.1%	36.7%	50.9%	8.9%	7,251
		(0.2, 0.5)	(2.7, 3.5)	(35.6, 37.9)	(49.8, 52.1)	(8.2, 9.5)	
	Very good	0.3%	2.1%	28.4%	55.3%	13.9%	964
		(0.1, 1.0)	(1.3, 3.2)	(25.7, 31.4)	(52.1, 58.4)	(11.9, 16.2)	
	n	104	825	6,896	6,260	1,003	15,088

Notes: Unweighted percentages using the balanced panel (respondents who participated in all three waves of the MxFLS and have complete SAH information); n=15,088. 95% confidence intervals are in parenthesis.
Source: Own estimates based on the Mexican Family Life Survey (MxFLS).

Figure 5 better depicts the behaviour of mobility in health over the period studied. While the point estimate of the Cowell and Flachaire mobility index (equations 10 to 12) indicates a decrease in mobility, the change is not statistically significant. This result holds for different values of the sensitivity parameter α.

Figure 5. Mobility in health during health insurance expansion in Mexico (balanced panel, weighted estimates)



Note: mobility is measured with the Cowell and Flachaire mobility index; 95 per cent confidence intervals, estimated using bootstrap with 1000 replications, are in brackets. The balanced sample includes the respondents who participated in all three waves of the MxFLS and have complete SAH information (proxy information considered); n=15,088.
Source: own estimates based on the Mexican Family Life Survey (MxFLS).

6.3. Robustness checks

This section examines whether some of the assumptions made to obtain the results in the previous section are likely fulfilled. In particular, we have a better look at the potential effects of non-response and the choice of the inequality measure.

6.3.1. Reconsidering non-response

To assess whether attrition may be biasing inequality estimates, we recalculated the Cowell and Flachaire index using the unbalanced sample. Table 5 shows that these estimates are consistent with the main results discussed above.[10] In sum, they suggest that health inequality increased between 2002 and 2009, although the changes are

[10] The sample used to estimate the weighted figures for waves 2 and 3 is slightly lower as some individuals have no weights assigned in the survey databases. While there is no clear explanation, the consistency of the unweighted results using the same sample suggests that the impact of these missing weights is negligible.

only statistically significant if the upward-looking status concept is adopted. This conclusion holds, however, for negative and positive values of the parameter α.

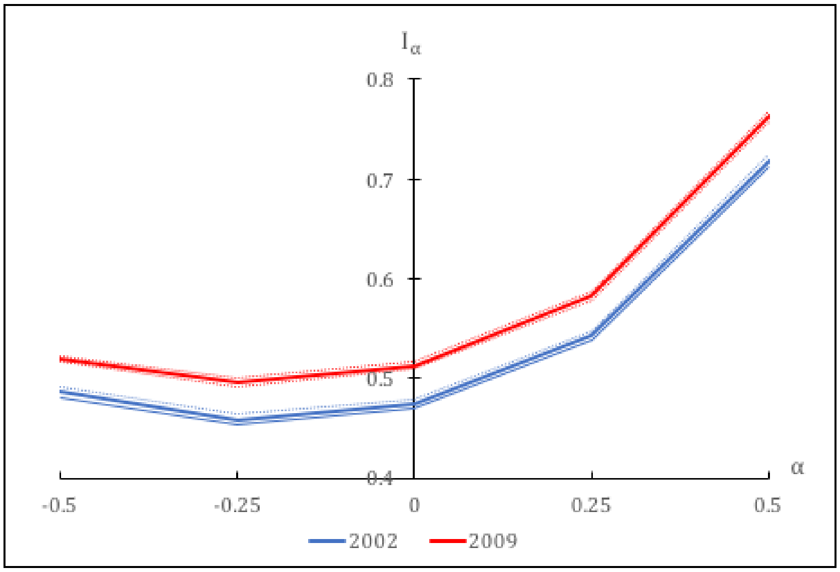
In addition, we recalculated the Cowell and Flachaire indices for the balanced sample without weights. While the results are again similar, these estimates provide stronger evidence of an increase in health inequality between 2002 and 2009, as this change is not only statistically significant for the upward looking definition of status but also for the downward looking definition (figures 6 and 7). The unweighted estimates of health mobility, on the other hand, confirm that it remained stable over the period studied. To facilitate the comparison, the first 2 columns of table 6 show the weighted estimates of mobility that correspond to figure 5 above, while the last two columns show the unweighted estimates.

Table 5. Health inequality during the public insurance expansion in Mexico. Unbalanced sample; weighted and unweighted estimates

α	2002 (n = 21,610)	2005 (n = 18,194)	2009 (n = 17,572)	2002 (n = 21,610)	2005 (n = 19,091)	2009 (n = 17,635)
	weighted results			unweighted results		
Panel A. Downward looking status						
-0.5	0.572	0.580	0.574	0.573	0.573	0.582
	[0.561, 0.582]	[0.566, 0.591]	[0.561, 0.585]	[0.567, 0.578]	[0.566, 0.579]	[0.576, 0.588]
0	0.506	0.516	0.512	0.510	0.512	0.519
	[0.496, 0.517]	[0.505, 0.526]	[0.498, 0.525]	[0.505, 0.514]	[0.507, 0.516]	[0.513, 0.524]
0.5	0.740	0.756	0.756	0.744	0.750	0.762
	[0.725, 0.756]	[0.740, 0.772]	[0.733, 0.774]	[0.739, 0.749]	[0.745, 0.755]	[0.756, 0.767]
Panel B. Upward looking status						
-0.5	0.492	0.505	0.519	0.492	0.494	0.518
	[0.484, 0.501]	[0.497, 0.512]	[0.509, 0.529]	[0.488, 0.496]	[0.491, 0.498]	[0.515, 0.521]
0	0.483	0.502	0.514	0.483	0.491	0.514
	[0.473, 0.493]	[0.492, 0.510]	[0.503, 0.525]	[0.479, 0.487]	[0.487, 0.496]	[0.510, 0.518]
0.5	0.729	0.752	0.764	0.730	0.740	0.766
	[0.713, 0.744]	[0.737, 0.768]	[0.745, 0.781]	[0.725, 0.735]	[0.733, 0.745]	[0.760, 0.771]

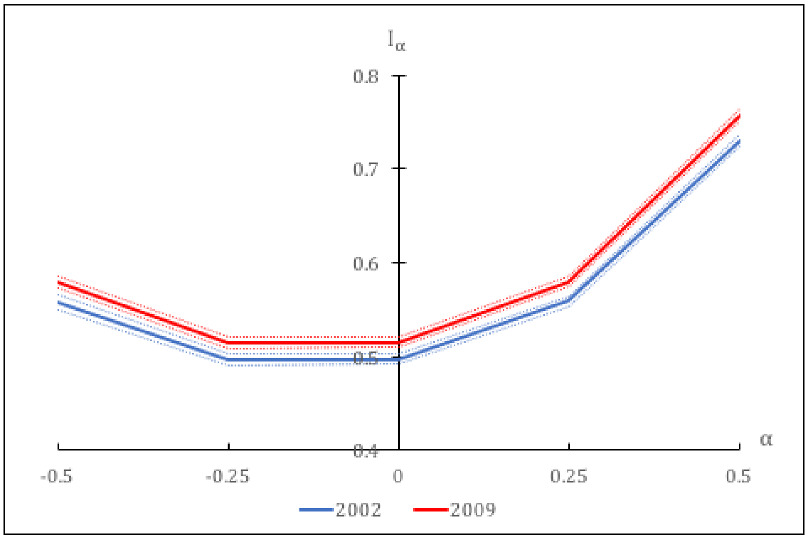
Notes: Inequality is measured with the Cowell and Flachaire inequality index; 95 per cent confidence intervals, estimated using bootstrap with 1000 replications, are in brackets. The unbalanced sample includes the respondents who participated in any of the three waves of the MxFLS (except for new entrants at wave 2 and 3) and have complete SAH information (proxy information considered).
Source: Own estimates based on the Mexican Family Life Survey (MxFLS).

Figure 6. Health inequality during the public insurance expansion in Mexico. Downward looking status (balanced sample, unweighted estimates)



Notes: Inequality is measured with the Cowell and Flachaire inequality index; the dotted lines represent 95 per cent confidence intervals estimated using bootstrap with 1000 replications. The balanced sample includes the respondents who participated in all three waves of the MxFLS and have complete SAH information (proxy information considered); n=15,088.
Source: Own estimates based on the Mexican Family Life Survey (MxFLS).

Figure 7. Health inequality during the public insurance expansion in Mexico. Upward looking status (balanced sample, unweighted estimates)



Notes: Inequality is measured with the Cowell and Flachaire inequality index; the dotted lines represent 95 per cent confidence intervals estimated using bootstrap with 1000 replications. The balanced sample includes the respondents who participated in all three waves of the MxFLS and have complete SAH information (proxy information considered); n=15,088.
Source: own estimates based on the Mexican Family Life Survey (MxFLS).

Table 6. Health mobility during the public insurance expansion in Mexico. Balanced sample, weighted and unweighted estimates

α	2002-2005	2005-2009	2002-2005	2005-2009
	weighted results		unweighted results	
-0.5	0.227	0.214	0.236	0.221
	[0.211, 0.243]	[0.198, 0.229]	[0.224, 0.247]	[0.211, 0.232]
0	0.164	0.155	0.171	0.162
	[0.156, 0.172]	[0.148, 0.163]	[0.166, 0.175]	[0.158, 0.167]
0.5	0.150	0.143	0.155	0.152
	[0.143, 0.157]	[0.138, 0.150]	[0.151, 0.159]	[0.148, 0.156]
1	0.163	0.160	0.167	0.170
	[0.156, 0.171]	[0.153, 0168]	[0.163, 0.172]	[0.164, 0.175]

Notes: Mobility is measured with the Cowell and Flachaire mobility index. 95 per cent confidence intervals estimated using bootstrap with 1000 replications are in brackets. The balanced sample includes the respondents who participated in all three waves of the MxFLS and have complete SAH information (proxy information considered); n=15,088.
Source: own estimates based on the Mexican Family Life Survey (MxFLS).

As noted before, proxy information of SAH was considered to avoid missing a large number of observations due to item non-response. If the individuals with proxy information, however, are systematically different from the rest of the sample, the results would be biased. The indices were therefore recalculated using only the information directly reported by the individual. Table 7 shows that the magnitude of these estimates is only slightly lower, but the pattern is the same. If we use the downward looking definition of status, no significant change is found between 2002 and 2005, but if we use the upward looking definition the increase in health inequality is statistically significant.

Table 7. Health inequality during the public insurance expansion in Mexico. Balanced sample with no proxy SAH information, weighted estimates

	2002	2005	2009
Panel A. Downward looking status			
α=0	0.483	0.499	0.491
	[0.468, 0.497]	[0.485, 0.511]	[0.475, 0.505]
Panel B. Upward looking status			
α=0	0.471	0.500	0.511
	[0.457, 0.485]	[0.488, 0.512]	[0.499, 0.522]

Notes: Inequality is measured with the Cowell and Flachaire inequality index; 95 per cent confidence intervals, estimated using bootstrap with 1000 replications, are in brackets. The balanced sample includes the respondents who participated in all three waves of the MxFLS and have complete SAH information (no proxy responses are considered for these estimates); n=11,897.
Source: own estimates based on the Mexican Family Life Survey (MxFLS).

Since attrition does not seem to affect the conclusions, we conducted a final test focused on item non-response. As mentioned above, imputation methods are widely used to deal with this type of non-response.[11] Survey administrators of the BHPS, for example, use hot-deck imputation and predictive mean matching, depending on the nature of the variable that is being imputed (Jenkins, 2011). But many different approaches are available. Here we use multiple imputation to account for uncertainty in the imputation strategy (Rubin 1987). In particular, we use multivariate imputation with chained equations (MICE) to take advantage of any SAH information available for individuals with missing values for some waves. This imputation method basically imputes multiple variables iteratively through a sequence of univariate imputation models, with fully conditional specifications of prediction equations. The model specified for the univariate imputation was an ordered logistic regression, with sex, age group, area and region of residence, marital

[11] Imputation methods can also be applied to replace the missing values caused by attrition (see an application imputation methods to deal with attrition in health surveys in Härkänen et al. 2016). As other simpler tests described above suggest that the effect of attrition is negligible in this case, however, we only use imputation to replace missing values of SAH for those who participated in all the three waves of the MxFLS.

status, education, household size, and participation in the labour market as independent variables.[12]

The sample for the imputed exercise include all the individuals who participated in all three waves of the MxFLS (n = 19,971; see table 1 above); proxy information was ignored, i.e., SAH responses provided by proxy informants were also treated as missing values. Since we also used percentile bootstrap with 1,000 replications to calculate confidence intervals, the number of imputations was set to five to simplify the computation procedure.[13] This implies that we estimated the inequality index for all five imputed datasets generated for each bootstrap sample. The estimated values of the index for each bootstrap sample were combined using Rubin's rule (1987), which basically amounts to calculating an average. Table 8 shows that the results obtained for the parameter α=0 are similar to those presented above. An increase in health inequality between 2009 and 2009 is noted, although the increase is statistically significant for both the downward and upward looking definition of status.

Table 8. Health inequality in Mexico. Balanced sample with multiple imputation of SAH, unweighted estimates

	2002	2005	2009
Panel A. Downward looking status			
α=0	0.498	0.506	0.512
	[0.494, 0.504]	[0.501, 0.512]	[0.506, 0.519]
Panel B. Upward looking status			
α=0	0.477	0.492	0.515
	[0.473, 0.481]	[0.487, 0.500]	[0.510, 0.521]

Notes: Inequality is measured with the Cowell and Flachaire index; 95 per cent confidence intervals, estimated using bootstrap with 1000 replications, are in brackets. The balanced sample includes the respondents who participated in all three waves of the MxFLS; n=19,896. Missing information of SAH was imputed.
Source: own estimates based on the Mexican Family Life Survey (MxFLS).

[12] Some of the independent variables had incomplete information for 75 individuals of the balanced panel (less than 0.5 per cent). These observations were excluded from the analyses.
[13] According to Schafer (1999) there is normally no practical benefit to using more than 5 imputations.

6.3.2. Measuring inequality with the Generalised Entropy index

Although the Generalised Entropy measures (GE_α) are suitable for cardinal variables, in this section we assess whether the results obtained using this indicator substantially vary from the results obtained using the Cowell and Flachaire index. The GE_α index was calculated using equations 5 to 7 above with the status of individual i, s_i, simply indicated by the category number of SAH (1 for very bad SAH, 2 for bad SAH, and so on).

Table 9 shows that the Generalised Entropy estimates are consistent with those obtained using the upward looking definition of status. For α=-1,0,1, this measure indicates that inequality in health increased over the period studied. These results hold for the balanced and unbalanced panel, with or without weights, except for some alphas for the weighted figures where the change between 2002 and 2009 is not statistically significant.

Costa-Font and Cowell (2016) had previously analysed the correlation between health inequality rankings across 70 countries using both the Cowell and Flachaire index (with the downward and upward looking definitions of status) and the GE index for different values of the sensitivity parameter α. Their results indicate that both measures resulted in similar patterns of inequality across countries only for the extreme case of $\alpha = 0.99$. This analysis, however, shows that both indices can give more consistent results for the analyses of within-country inequality patterns.

Table 9. Health inequality in Mexico using the Generalised Entropy Index

α	2002	2005	2009	2002	2005	2009
	weighted results			unweighted results		
Panel A. Balanced panel						
-1	0.024	0.023	0.026	0.024	0.022	0.028
	[0.022, 0.025]	[0.022, 0.025]	[0.025, 0.028]	[0.023, 0.025]	[0.021, 0.023]	[0.027, 0.029]
0	0.021	0.021	0.023	0.021	0.020	0.025
	[0.020, 0.022]	[0.020, 0.022]	[0.022, 0.024]	[0.020, 0.022]	[0.020, 0.021]	[0.024, 0.026]
1	0.019	0.020	0.022	0.020	0.019	0.023
	[0.019, 0.020]	[0.019, 0.021]	[0.021, 0.023]	[0.019, 0.020]	[0.019, 0.020]	[0.022, 0.024]
n	15,088	15,088	15,088	15,088	15,088	15,088
Panel B. Unbalanced panel						
-1	0.025	0.026	0.027	0.025	0.024	0.028
	[0.024, 0.027]	[0.024, 0.028]	[0.026, 0.029]	[0.024, 0.026]	[0.023, 0.025]	[0.027, 0.030]
0	0.022	0.023	0.024	0.022	0.022	0.025
	[0.021, 0.023]	[0.022, 0.024]	[0.023, 0.025]	[0.022, 0.023]	[0.021, 0.022]	[0.024, 0.026]
1	0.021	0.021	0.022	0.021	0.020	0.023
	[0.020, 0.021]	[0.019, 0.021]	[0.022, 0.023]	[0.020, 0.021]	[0.020, 0.021]	[0.022, 0.023]
n	21,610	18,194	17,572	21,610	19,091	17,635

Notes: 95 per cent confidence intervals, estimated using bootstrap with 1000 replications, are in brackets. The balanced sample (panel A) includes the respondents who participated in all three waves of the MxFLS and have complete SAH information (proxy information considered); the unbalanced sample (panel B) includes the respondents who participated in any of the three waves of the MxFLS (except for new entrants at wave 2 and 3) and have complete SAH information.
Source: own estimates based on the Mexican Family Life Survey (MxFLS).

7. Discussion

This paper has empirically examined the change in pure health inequality and mobility in Mexico after the introduction of a major public insurance expansion that universalise access to health care, so-called Seguro Popular. Rather than examining socio-economic inequalities in health, and more specifically a concentration index of health on income, we used a class of measures appropriate to deal with categorical indicators of SAH to analyse pure health inequalities and mobility.

We find that, the distribution of health worsened in Mexico between 2002 and 2009, although the direction of change is only consistent for the upward looking definition of status. Together with the lack of mobility in health observed, we may thus conclude that Mexico is becoming more rigid over time insofar as the distribution of health status. While short study periods could be expected to provide little opportunity for movement in general, Hauck and Rice (2004) actually found evidence of large mobility in mental health in the UK over the 1990-2000 decade. In contrast, Contoyannis *et al.* (2004) found strong persistence in self-reported health status in the UK in the same period. Our findings are in line with the latter.

Our results complement previous findings on the effect of insurance expansion on health status. Teruel et al. (2012) previously analysed the effects of increased coverage through the Seguro Popular on perceived health status. They used data from the MxFLS and propensity score matching to create a suitable comparison group drawn from those still uninsured at the time of collection of the third wave. At baseline, those who gained insurance through the SP were more likely to report bad health than the comparison group, but the analysis showed that a 6 per cent increase in the probability of reporting good health among the former can be attributed to the programme. How can we reconcile this result with ours? While the Seguro Popular may helped improve SAH among beneficiaries, it seems that other factors shape the overall distribution of health.

Unfortunately, the limited variability in the available data for Mexico do not allow analysing the extent to which different economic, institutional, and environmental factors affect health disparities alongside income inequality influence the emergence of health disparities. In particular, Baeten et al. (2013) argue that the contribution of income inequality to health inequality is between 25 per cent and 30 per cent. Wang and Yu (2016) also show that common indicators of income inequality such as the Gini coefficient and the Theil index are positively associated with health disparities. Income inequality in Mexico declined over the past decade, however (Esquivel 2015, OCDE 2014). This decline has been attributed to increases in remittances among low income households, and reductions in labour income and non-labour income (government transfers) inequalities (Esquivel 2015, Esquivel et al. 2010). Costa-Font and Cowell (2016), on the other hand, suggest that institutional performance, in particular better government effectiveness, is associated with health inequality declines. According to the Worldwide Governance Indicators (Kaufmann et al. 2010), government effectiveness in Mexico declined from 0.24 in 2002 to 0.17 in 2009. Other indicators of governance such as regulatory quality, control of corruption, and political stability and absence of violence present much larger drops. Therefore, these factors could be key to explain the pattern of health disparities in Mexico. Lifestyle indicators should be considered too. Specifically, Mexico has been subject to an obesity epidemic in the period (e.g. Colchero and Sosa-Rubí 2012), which has affected more deprived population groups that might have benefited from the SP.

From a policy perspective, if the distribution rather than just overall levels of health is indeed a concern, a first step should be to start monitoring health inequality using measures such as those introduced in this study. While international organisations such as the WHO and the OECD normally include Mexico in their endeavour to monitor inequality in health (e.g WHO 2000, OECD 2014, 2015), there is no clear initiative at the national level. The National Council for the Evaluation of Social Development Policy (Coneval), the institution in charge of poverty measurement and other activities oriented towards the achievement of social development objective, currently estimates some inequality indicators, but these only include the Gini coefficient and two inter-decile ratios to measure income disparities. Furthermore, there is limited coordination between public health and health care system initiatives, which can explain why measures of health equity do not show major shifts since the introduction of the SP.

Wwhile further analysis on the potential drivers of health inequalities is needed, the Mexican experience suggests that insurance coverage can improve health levels but may be not enough to reduce health disparities and promote health mobility. Indeed, health inequality and mobility likely depend on many factors beyond health care.

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