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Does publicly provided health care affect migration? Evidence from Mexico
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Abstract

This paper investigates whether social policies affect migration propensity. I exploit the random expansion of a publicly provided healthcare programme in Mexico, as well as the panel dimension and the timing of the Mexican Family Life Survey. Difference-in-differences estimations reveal that non-contributory healthcare increases internal migration by freeing up care (time) constraints and strengthening household economic resilience in the face of health-related shocks. However, the alleviation of financial and time constraints is not significant enough to alter international migration, which is more costly by nature. Results point to the relevance of including both resident and non-resident household members in assessing the effects of social policies on labour market behaviours. They suggest that, in the setting studied, publicly provided healthcare complements, rather than substitutes for, informal livelihood strategies. Relaxing binding financial and time constraints enables labour force detachment of working-age members in affiliated households.

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1 Introduction

Publicly provided healthcare has received growing attention in the academic and policy debate. Extending public healthcare is claimed to be one of the most effective ways of reaching universal health coverage (Jamison et al., 2013). Positive impacts on health and poverty are more and more acknowledged.[1] Still, concerns about adverse labour market effects, such as disincentives to work or switching from formal to informal work, have been raised,[2] and empirical evidence on labour market outcomes overall remains mixed.[3]

By providing (near-)poor households with the means to deal with risk, not only might (formal) social protection programmes affect labour market behaviours, they might simultaneously alter household livelihood strategies. However, relationships between social protection and livelihood strategies have not attracted much scrutiny. Since Stark and Bloom's (1985) New Economics of Labour Migration (NELM), one of these, migrating, has been seen as a coping strategy – a means for households to respond to shocks (Sabates-Wheeler and Waite, 2003). In this sense, accessing different sources of living through welfare could affect the need to migrate (Hagen-Zanker and Leon-Himmelstine, 2013). Providing healthcare for free could directly affect migration by relaxing financial constraints, reducing the occurrence and duration of health shocks, and therefore enabling (healthier) working-age household members to migrate. It could indirectly influence migration through its effects on local labour markets by inducing them to reallocate time from care giving to working outside their households, for instance prompting their entry into the labour market, which might involve migrating.

In analysing the dynamics between publicly provided healthcare and labour force attachment of working-age members in affiliated households, this paper aims at shedding light on the mixed evidence on the links between publicly provided healthcare – welfare programmes in general – and labour market behaviours, often found in the literature. Because of their very focus on recipients and household members who live with them, considering spatial mobility by looking at implications for migration, that is including non-resident household members and/or household members who (might have) migrated in the analysis, could explain the heterogeneity in existing results. Household affiliation to healthcare might enable working-age beneficiaries to migrate, who would then 'disappear' from estimation samples. Not accounting for the potential effects of social protection programmes on the likelihood to migrate might question the reliability of results obtained for labour market outcomes.

Moreover, in studying specifically whether publicly provided healthcare helps surmounting financial and care constraints, this paper clarifies the migration decision-making process as well as livelihood strategies. Given the documented effects of migration on development,[4] adding evidence on whether migration complements or substitutes for social policies could give insights into what prevents migration. This is necessary to improve the design and target of policies seeking to remove impediments to mobility for those who could benefit from welfare gains, and to leverage the contribution of migration to the development of origin communities (Cazzuffi and Modrego, 2017).

To assess how migration relates to publicly provided healthcare, this paper exploits the random expansion of the programme *Sistema de Protección Social en Salud*, introduced in Mexico in 2003, that aims at improving economic resilience through access to health services (almost) for free. Offered to some 50 million Mexicans without social security, it institutionalised a pilot programme, *Seguro Popular de Salud*, running from October 2002 to December 2003, that has gradually expanded across Mexico. This paper takes advantage of the timing

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[1] See for instance Bitler and Hoynes (2016) on the mitigating effect of safety nets on poverty in times of crisis, Bitler et al. (2017) on child poverty and Bitler and Hoynes (2013) on immigrant populations, or Bitler and Hoynes (2008) on long-run health outcomes.

[2] See for instance Cutler et al. (1996), Gruber and Simon (2008) or Levy and Schady (2013).

[3] For a review, see Ravallion (2003); for evidence from the United States (US), Baicker et al. (2014), Dave et al. (2015) or Garthwaite et al. (2014); for evidence from Mexico, Azuara and Marinescu (2013), Bosch and Campos-Vázquez (2014) or del Valle (2016).

[4] On the positive development impacts of migration, see for instance Adams and Page (2005); on its negative effects, see Portes (2006) or Vullnerati and King (2008).

of the Mexican Life Satisfaction Survey (MxFLS), a three-wave household panel conducted from 2002, before the start of *Seguro Popular* pilot phase, to 2009. A difference-in-differences specification is used to compare changes in migration between individuals living in municipalities where this programme was introduced earlier and individuals living in municipalities where it was implemented later. The panel structure of the MxFLS allows controlling for endogenous migration by assigning changes in coverage to individuals based on their municipality of residence at the beginning of each time period. This paper focuses on short-run effects on internal and international migration. Effects on health status, time use, health-related shocks and credit worthiness are also taken into account.

Exposure to a change in coverage is found to raise internal migration, but to have a statistically insignificant effect on international migration. Estimates suggest that access to publicly provided healthcare might play a role in reducing credit and care constraints, enabling working-age household members to migrate in families vulnerable to adverse shocks. Accessing publicly provided healthcare appears to free up caregivers' time and strengthen household economic resilience in the face of health-related shocks. In contrast to contributory healthcare that tends to crowd out migration by tying affiliates to formal employment and so a specific location, non-contributory healthcare seems to complement, rather than substitute for, alternative livelihood strategies. Placebo tests confirm the validity of the identification strategy against threats of time-trending unobservables, and suggest that changes in migration propensity prior to the programme were negatively correlated with its expansion.

The rest of this paper is structured as follows. Section 2 describes how access to free healthcare could affect migration, section 3, Mexico's health insurance system and section 4, data and estimation strategy. Section 5 presents results. Section 6 concludes.

2   Publicly provided healthcare and migration

Departing from neo-classical models,[5] later theoretical approaches, such as the NELM, consider migration as a decision made within a household to maximise wealth, diversify income sources between household members and across space, and minimise risks (Stark and Bloom, 1985). In this regard, migrating can be viewed as an informal livelihood coping strategy (Sabates-Wheeler and Waite, 2003). Whether and which household members migrate is decided jointly by those likely to migrate and those likely to stay behind and to support migrants or benefit from their migration, through the receipt of remittances for instance. The decision to migrate depends on opportunities inside and outside households' places of residency, costs induced by moving and being absent, number and share of household dependants, and credit constraints. In other words, deciding to migrate is determined by financial and care (time) constraints.

Assuming that household members share their resources, accessing alternative sources of livelihoods thanks to healthcare (or other safety nets) could influence the decision to migrate directly and/or indirectly through effects on labour market behaviours. Figure 1 depicts these dynamics. By reducing the occurrence and duration of health shocks and health-related expenditures, publicly provided healthcare is expected to improve beneficiaries' health statuses, increase their disposable income and limit working-age members' time dedicated to caring for (potentially sick) dependants.

By minimising health-related shocks and expenditures, publicly provided healthcare might reduce the need to diversify income sources, i.e. working outside one's household and spatially reallocating its labour force. In a cost-benefit analysis, labour, migration and safety nets would be substitutes in this case. Accessing healthcare would crowd out work and migration, since households and individuals would not have to rely on the labour market and migration to minimise risks. In addition, the opportunity costs of migrating would increase, if it implies losing health insurance coverage. The alleviation of financial constraints would induce working-age members to work less or not to work, and hence, not to migrate.

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[5] See, for instance, Harris and Todaro's (1970) model of rural-urban wage differences.

On the other hand, non-contributory healthcare could support affiliated households and their working-age members in looking for work, locally or outside their communities of origin. This could be effected by improving health and enabling the reallocation of their labour force and financial resources, previously used to tackle household (dependant) and individual health shocks, towards remunerative activities outside their households. Healthcare, labour and migration would be complements. This would be particularly relevant in a context of imperfect labour markets where it is difficult and costly to hire non-household members to care for dependants.

First, by specifically targetting the incidence and dissemination of communicable diseases (Knaul et al., 2006), Pfutze (2015) and Conti and Ginja (2016) showed that the introduction of *Seguro Popular* has led to significant improvements in healthcare use and health status in Mexico. Evidence from a randomized experiment in Kenya also reveals that health insurance limits stress and cortisol levels (Haushofer et al., 2017). After investigating several channels, the authors conclude that, akin to a 'peace of mind' effect, the simple fact of having coverage improves sleep, in particular among more vulnerable people. That affiliation to healthcare improves physical and mental health is likely to enhance the productivity of working-age affiliates, which could strengthen their ability to migrate, consistent with a 'healthy emigrant effect'.[6]

Second, if financial constraints hinder the capacity of households to send migrants away, access to healthcare, i.e. a punctual but exogenous source of income, could relieve such constraints and alter the degree of labour attachment of working-age members in affiliated households (Hagen-Zanker and Leon-Himmelstine, 2013). By limiting daily and catastrophic health expenditures, and thus increasing disposable income, non-contributory healthcare could not only boost consumption, but also be used to directly finance migration. Bryan et al. (2014) show that, when households were randomly assigned a financial incentive in rural Bangladesh, 22% of recipient families sent a member away during the pre-harvest lean season. By limiting risks induced by migration, this cash incentive contributed to diversify income sources through migration.

Non-contributory healthcare could also be used to finance migration indirectly. Welfare programme entitlement has been shown to relax binding financial constraints. For instance, Angelucci (2015) shows that poor households' entitlement to an exogenous source of income through *Oportunidades*[7] increased emigration to the United States (US). Although cash transfers were mainly consumed, families who could not previously afford to migrate, used entitlement to *Oportunidades* as collateral to ask for loans and finance migration. By enabling households to overcome financial constraints, access to healthcare could directly and/or indirectly help households by supporting working-age members in looking for work, which might involve migrating. This is in line with evidence on other non-contributory programmes such as the Old Age Grant, a pension scheme targeting the elderly in South Africa.[8]

If there is any effect, effects might differ depending on household composition, as suggested by del Valle (2016). Because women tend to bear the greatest care giving burden in Mexican households, *Seguro Popular* might have differential impacts across gender. Affiliation might affect women's entry into the labour market in source communities rather than their propensity to migrate. Since women are more likely to take care of dependants than men, the expansion of health coverage might push women to enter the (informal) labour market, as shown by del Valle (2016),[9] now freed from caring for their dependants. The associated increase in disposable income and decrease in time constraints might not be significant enough to prompt them to migrate. In contrast, men, less likely to take care of dependants compared to women, might show greater

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[6] Migrating often involves many obstacles. The literature suggests that only the fittest, and hence healthiest, would successfully emigrate. The harder such obstacles, the stronger the positive health selection (Jasso et al., 2004).

[7] A Mexican conditional cash transfer (CCT) programme giving cash to households with children, conditional on children regularly attending school, visiting health clinics and looking after their diet.

[8] Inder and Maitra (2004), Posel et al. (2006), Sienaert (2008) and Ardington et al. (2009) find that non-contributory pension programmes induced migration within multi-generational households by alleviating budget constraints, whereby enabling working-age members to search for (and eventually find) work in urban areas.

[9] Since access to public healthcare is conditional on not working in the formal sector, *Seguro Popular* could affect the decision to work in the formal or informal sector by increasing the non-monetary benefits of informal work (Bosch and Campos-Vázquez, 2014; Azuara and Marinescu, 2013) and the relative price of formal (health) insurance (Gruber and Madrian, 1994). Moreover, this could encourage formal workers with employer-based health insurance to become informal with publicly provided, cheaper if not free, health insurance, and dissuade informal workers from becoming formal.

labour attachment flexibility. Upon affiliation to healthcare coverage, they might be more likely to leave source households in order to further diversify household income sources, as gender-differentiated evidence from South Africa shows.[10] Women would simultaneously work (part-time) outside of their households and take care of household dependants, as del Valle (2016) points out.

Moreover, since international migration tends to be more costly than internal migration, and as the entitlement to *Seguro Popular* does not represent an exogenous stream of income as such, but rather an increase in disposable income due to a reduction in health expenditures, this insurance might be more likely to affect internal than international migration. Internal and international migration bear different costs. Internal migration is less expensive and less risky, since conditions to migrate internally are easier to meet (Stecklov et al., 2005). International migration might only be affected when it is not more costly than internal migration, which is unlikely to be the case. Existing empirical findings suggest that safety nets have different, often opposite effects on domestic and global migration. While Chau et al. (2012), Inder and Maitra (2004), Posel et al. (2006), Sienaert (2008) and Ardington et al. (2009) have confirmed a positive effect of safety nets on internal migration via its funding channel, Angelucci (2015) does not find a consistent effect of *Oportunidades* on internal migration.

Figure 1: Linkages between publicly provided healthcare and migration
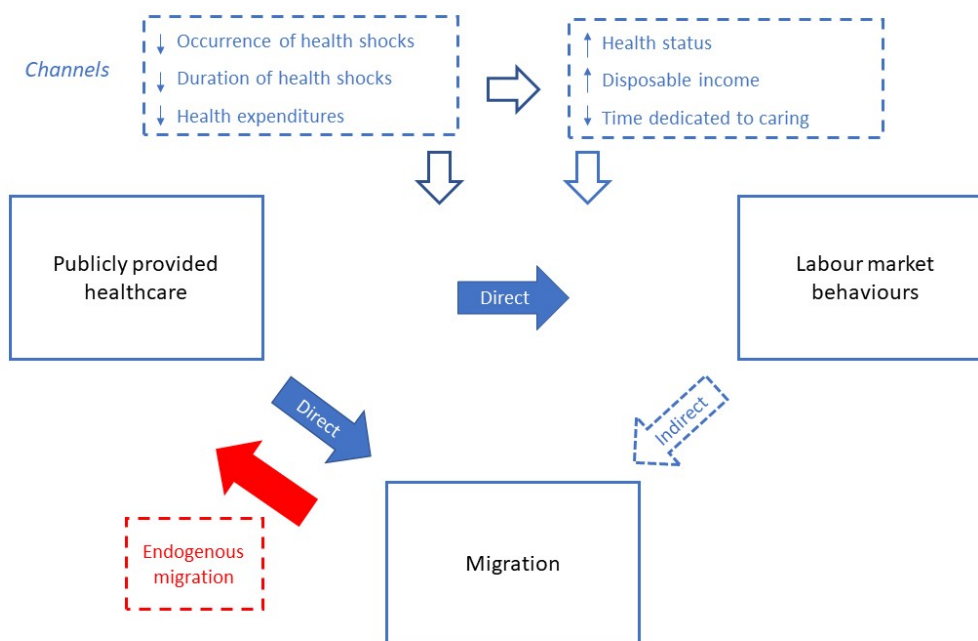
3   Mexico's health insurance system

Mexico's health system is divided into two sectors. Health services are provided upon contributions to social security institutions run by the government – mainly the *Instituto Mexicano del Seguro Social* (IMSS, Mexican Social Security Institute) in the private sector and the *Instituto de Seguridad y Servicios Sociales de los Trabajadores del Estado* (ISSSTE, Social Security and Services Institute for Public Workers) in the public sector. They are accessed through formally registered employers. Those out of the labour force or

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[10] Evidence from South Africa indicates that, when women are those affiliated, other family members, in particular men, tend to migrate (Sienaert, 2008).

working in the informal sector – about half the population – access a small number of underfunded services through the *Secretaria de Salud* (SS, Ministry of Health). As a result, there have been major gaps in resource allocation and inequalities between beneficiaries of these two healthcare systems (Frenk et al., 2009).

For this reason, the *Sistema de Protección Social en Salud* was introduced in 2003 to improve financial strength through health services. Offered to some 50 million individuals without social security, it guarantees subsidised, publicly provided basic universal healthcare services – essential primary and secondary services and high-complexity healthcare interventions, medication, laboratory and cabinet analyses. Affiliates have access to medical-surgical, pharmaceutical and hospital services to meet their health needs. It currently offers coverage to 275 medical interventions, as well as services that support people who do not have access to formal social security, and who suffer from high-cost diseases that can put their lives and family assets at risk. This programme has eliminated fees for services and drugs, removed access to health services and extended the supply of health services through medical infrastructure.

Healthcare is provided almost for free,[11] only to those not covered by any social security institution. To be affiliated, individuals must reside in Mexico and may not benefit from any other social security institutions. They can apply for (nuclear) family affiliation on a voluntary basis,[12] and have to provide the necessary information for a socio-economic evaluation of the family.[13] If an affiliate happens to be outside her place of residence, she is covered, as long as her policy is in force and she can refer to her entity of origin, for medical emergencies and in case of patients in transit.[14]

Seguro Popular was implemented in stages. It was introduced in October 2002 with a pilot phase in five states. 14 additional states adopted *Seguro Popular* until December 2003. As shown by Figure 2, expansion was relatively fast, with almost full coverage reached in 2011 (Pfutze, 2015). Although the roll-out of the programme in principle gave priority to the poorest areas with sufficient healthcare infrastructure, political and logistical considerations might have played a role (Barros, 2008; Díaz-Cayeros et al., 2006). Governors decided when to participate, and had some degree of autonomy in choosing when the programme would be implemented in eligible municipalities. However, others have concluded that the introduction of *Seguro Popular* was close to random with regard to *ex-ante* (2000) covariates at the municipality and state levels such as income, number of uninsured, industrial structure, informality and/or labour market outcomes (Aterido et al., 2010; Azuara and Marinescu, 2013; Bosch and Campos-Vázquez, 2014; del Valle, 2016).

Following these studies, this paper exploits the random expansion of *Seguro Popular* at the municipality level to examine its effect on migration propensity. As in del Valle (2016), it is argued that random variation in roll-out can be identified by focusing on municipalities where *Seguro Popular* was introduced in its expansion phase – in the middle of its roll-out – from 2004 onwards. Municipalities that experienced a significant change in coverage from 2002 to 2004, i.e. in its pilot phase, are excluded. Following Alcaraz et al. (2016), treated municipalities are defined as municipalities where the change in coverage from the last quarter of 2004 to 2008 was at least 10% or greater. Control municipalities are those where the change in coverage was strictly less than 10% in both the 2002-2004 and 2004-2008 periods. Municipalities where the change in coverage was at least 10%, but strictly less than 20%, are excluded. This is done to ensure estimates are not biased by treatment contagion (Alcaraz et al., 2016). This would be the case if control municipalities were similar to treated municipalities that experienced a relatively small increase in coverage that was big enough to be defined as treated.

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[11] For a detailed review of this initiative, see Bosch et al. (2012).

[12] That is (i) spouses or cohabiting partners, (ii) parents who are not married, (iii) children younger than 18 years old, (iv) minors who are part of the household and are related to spouses by blood, partners or parents, (v) single children up to 25 years old who can prove to be students, (vi) dependant disabled children, (vii) straight-line direct ancestors over 64 years of age, who are economic dependants and live in the same household, and (viii) persons not related to spouses, partners or parents, but who live in the same house and depend economically on it and are under 18 or disabled dependant of any age.

[13] Such as proof of address, a unique code of population registry, birth certificate, official identification with photograph of the person who acts as family head, receipt of payment for the corresponding family fee, except in the case of households that enter the non-contributory regime due to their socio-economic status.

[14] To receive the services offered by *Seguro Popular* in a hospital or health center in another state, an affiliate must submit eiher a voting card, a Mexican passport or a military service primer.

Figure 2: Seguro Popular beneficiaries and implementation phases
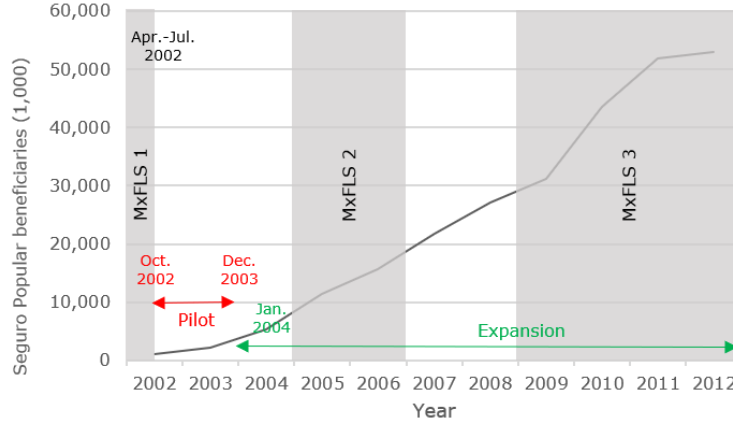
4 Data and empirical strategy

4.1 Empirical strategy

To estimate the effect of healthcare provision on migration, this paper uses a difference-in-differences estimator, by comparing the difference in migration propensity between treated and control municipalities before and after the expansion of *Seguro Popular*. It is assumed that the difference in migration propensity would have remained the same had *Seguro Popular* not been introduced. Moreover, municipality fixed effects take into account unobserved characteristics at the treatment (municipality) level. Any unobserved variable that might be related to migration is assumed to be uncorrelated with *Seguro Popular* expansion, conditional on observed covariates. Estimate robustness to this assumption is assessed later. Treated municipalities with a change in coverage of at least 10% and strictly less than 20% are excluded to avoid treatment contagion, as explained in section 3.

Regression (1) estimates the impact of living in a municipality m at time t on the propensity to have migrated in the subsequent wave of the survey. A linear probability model with municipality fixed effects is run as shown below:

$$Y_{imt} = \alpha + \beta_1 Treated_m + \beta_2 Post_t + \beta_3 (Treated * Post)_{mt} + \delta' X_{imt} + \gamma Municipality_m + u_{imt} \quad (1)$$

where Y represents the outcome variable of interest, between wave migration, of a 21-65 year-old individual successfully interviewed in at least two consecutive waves, living in municipality m at the beginning of time period t. $Treated$ is a binary variable taking value 1 if changes in *Seguro Popular* coverage in a municipality m in which respondent i lived at the beginning of each time period were strictly smaller than 10% between the last quarter of 2002 and the last quarter of 2004, and of at least 10% between the last quarter of 2004 and the last quarter of 2008 (treated municipalities); 0, if changes in coverage in a municipality were strictly less than 10% in both periods (control municipalities). $Post$ and $Municipality$ are respectively time and municipality fixed effects. β_3 is the difference-in-differences estimator.

X is a vector of individual, household and municipality characteristics. It includes gender, age, years of schooling of respondents; gender, age, years of schooling, indigenous origins of household heads; household dependency ratios for 0-7, 8-14, 15-20 and 66 years old and more; whether a household has experienced any economic shock in the preceding five years; household wealth (asset) index, excluding farm-related

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assets; and lagged indices of marginalization at the municipality level. Compiled by *Consejo Nacional de Población* (CONAPO), marginalization indices are based on several indicators of education, dwelling and income, collected every five years, to inform about the degree of poverty, inequality and exclusion at some administrative level. The higher marginalization indices are, the poorer localities are. Furthermore, this variable should control for the presence of welfare programmes like *Oportunidades* as their introductions are based on such marginalization indices. u_{imt} is the error term. Standard errors are robust to heteroscedasticity.

Variations in cross-municipality (-state) welfare benefit generosity could incite low-income families to migrate, if they were not provided with similar benefits in their current places of residence. Existing findings point to an increase in mobility from low- to high-benefit areas (Moffitt, 1992). As Moffitt (1992) explains, the literature on migration effects, rather scarce, is limited by econometric and methodological issues, in particular in the use of a truly random source of variation in welfare schemes. If individuals might (have) migrate(d) to municipalities where *Seguro Popular* was introduced to access healthcare, any migration observed between waves might be the result of individuals pulled to migrate to benefit from the programme, rather than *Seguro Popular* affiliation enabling them to migrate. This paper takes advantage of the timing and the panel dimension of the MxFLS to control for this potential source of endogeneity (reverse causality). Changes in *Seguro Popular* coverage are assigned to respondents based on their municipality of residence at the beginning of each period – their 2002 municipality of residence when looking at the relation between changes in coverage and migration between MxFLS wave 1 (2002) and 2 (2005), and their 2005 municipality of residence when looking at the relation between changes in coverage and migration between MxFLS wave 2 and 3 (2009).

4.2 Data

Data source

This analysis exploits the timing of the MxFLS, a three-wave household panel conducted from April to July 2002, before the start of *Seguro Popular* pilot phase, in 2005-2006 and in 2009-2012. The MxFLS specifically identifies all migrants, internal or international, even those who permanently moved to the US. This avoids potential biases of other data sets used in the literature, such as the undercount of Mexican migrants to the US as in Chiquiar and Hanson (2005), or migration information based on recall as in Orrenius and Zavodny (2005) (Kaestner and Malamud, 2014). In addition, the MxFLS provides individual and household level details on demographic and socio-economic characteristics.

Information on individual and household administrative records of *Seguro Popular* by municipality by quarter comes from Bosch and Campos-Vázquez (2014). Of the number of affiliates to *Seguro Popular* in each quarter from 2002 to 2009, municipalities that experienced a significant change in coverage, at least 10%, from the last quarter of 2002 to the last quarter of 2004 are excluded to simultaneously take advantage of the timing of the MxFLS and the random variation in the middle of *Seguro Popular* implementation, from 2004 onwards. As explained in section 3, this paper follows Alcaraz et al. (2016) in defining treated municipalities as municipalities where the change in coverage was at least 10% or greater from the last quarter of 2004 to the last quarter of 2008. Control municipalities are those where changes in coverage were less than 10% in both the 2002-2004 and 2004-2008 periods.[15] Individuals living in municipalities where the change in coverage was at least 10% but less than 20% are excluded, not to bias estimates because of treatment contagion (Alcaraz et al., 2016).

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[15] The last quarters of 2004 and 2008, rather than 2005 and 2009, are selected, since individuals are not expected to change their behaviour instantaneously due to changes in coverage.

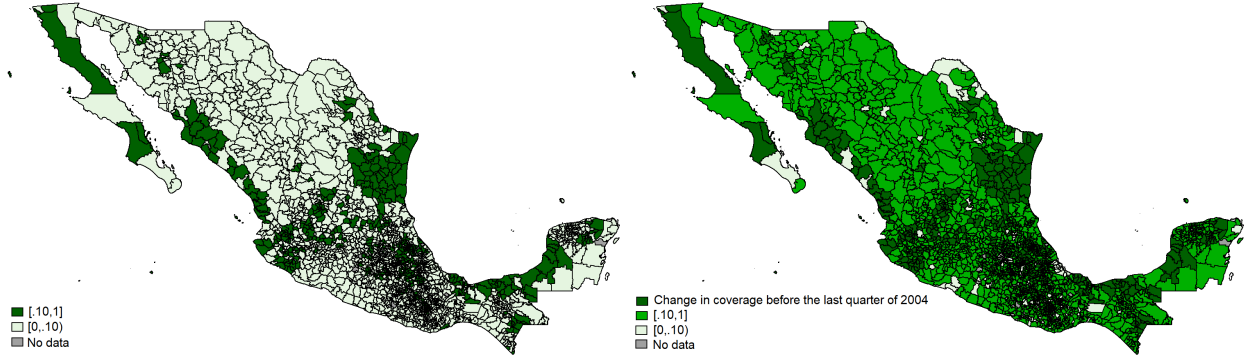Figure 3: Municipalities by change in coverage rate in the last quarter of 2004 (l) and 2008 (r)

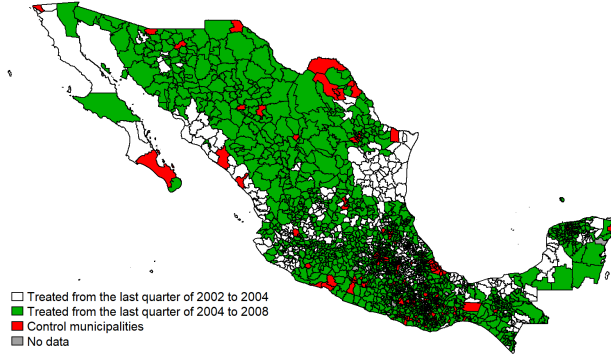Figure 4: Municipalities by treatment status

Figure 3 illustrates Mexican municipalities by changes in *Seguro Popular* coverage rate up to the last quarter of 2004 (pre-treatment period) on the left, and from the last quarter of 2004 to the last quarter of 2008 (treatment period) on the right.[16] Figure 4 represents municipalities included in this analysis by treatment status.

Descriptive statistics

As in Kaestner and Malamud (2014), the estimation sample is limited to 21-65 year-old men and women, successfully interviewed in at least two waves of the survey. After dropping observations with missing variables, a sample of 5,872 unique respondents interviewed across two time periods is obtained, forming an unbalanced panel of 9,431 observations. The dependant variable is a binary variable taking value 1 if an individual has migrated between two subsequent waves, between MxFLS wave 1 and 2, and/or between MxFLS wave 2 and 3; 0, otherwise.

Descriptive statistics provided in Tables 1 and 2 reveal differences across treatment and control groups. Individuals living in treated municipalities are relatively older, and have a lower level of education than individuals living in control areas. They come from bigger and poorer households that are more likely to have experienced an economic shock in the previous 12 months. Their household heads are also older, more likely to be men, with lower education, and more likely to be from an ethnic minority background. Lagged marginalization indices of treated municipalities are higher than in control municipalities. These differences are observed in both the 2002-2005 and 2005-2009 periods, except regarding the age of household

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[16] On this part of Figure 3, municipalities that experienced a significant change in coverage rate by the last quarter of 2004 are coloured in dark green regardless of their change in coverage rate in the second time period.

heads and the occurrence of household economic shocks. The analysis takes into account such differences and unobserved heterogeneity by including time fixed effects, alternatively municipality and individual fixed effects, and time-varying characteristics.

Table 1: Descriptive statistics of estimation sample

	Full sample Mean	SD	Treated Mean	Control Mean	t-test
Outcome variables (period end)					
Migrated between waves	0.063	0.244	.052	.074	-4.49***
Internal	0.058	0.233	.044	.071	-5.64***
International	0.008	0.086	.010	.005	3.01***
Control variables (period start)					
Age	39.991	11.515	40.609	39.411	5.05***
Male	0.468	0.499	.472	.464	0.72
Years of schooling	6.894	4.073	5.531	8.174	-33.29***
Household head					
Age	47.739	13.105	48.324	47.19	4.21***
Male	0.832	0.374	.857	.808	6.39***
Years of schooling	6.176	4.199	4.956	7.319	-28.45***
Indigenous	0.147	0.354	.221	.077	20.11***
Household dependency ratio					
0-7	0.131	0.164	.137	.125	3.60***
8-14	0.142	0.166	.152	.133	5.46***
15-20	0.109	0.148	.117	.102	4.93***
66 and more	0.036	0.105	.039	.033	2.93***
Household economic shock	0.290	0.454	.309	.272	3.98***
Household wealth index	0.140	1.425	-.306	.559	-30.91***
Lagged marginalization index	-0.983	0.889	-.344	-1.582	94.12***
Observations	9,431		4,565	4,866	

Notes: Means and standard deviations (SD) of variables of interest of the estimation sample, 5,872 unique individuals aged 21-65 years old, forming an unbalanced panel of 9,431 respondents interviewed across 2002-2009. Means of *health status*, *spent on health* and *health expenditures* variables are respectively based on 8,860, 9,183 and 9,099 observations. *** p<0.01, ** p<0.05, * p<0.1.

While individuals in treated municipalities have, on average, a lower probability to migrate between waves than individuals in control municipalities, Table 2 indicates that individuals living in treated municipalities display a significantly lower propensity to migrate in the pre-treatment period, but not any significant difference in the post-treatment period. Migration propensity before the introduction of the programme was lower in treated areas than in control areas. In contrast, migration propensity in control areas slightly decreased between the two periods. This downward shift in control municipalities is consistent with migration trends in Mexico. While the proportion of the Mexican population who lived in a state different from their state of birth increased from 10.6% in 1940 (2,081,000 people) to 19.2% in 2000 (18,752,000), it remained almost constant in relative terms in 2010 (19.3%, 13,976,000 people) (Pimienta-Lastra et al., 2012). These downward trends are confirmed for internal migration between functional territories (Cazzuffi and Pereira-López, 2016).[17] International migration has also been decreasing, likely because of a labour-foreign direct investment (FDI) effect in receiving Mexican states (Aroca and Maloney, 2005), combined with increasing costs involved in international migration (Orrenius, 2001). Descriptive statistics of the estimation sample suggest that the expansion of healthcare coverage

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[17] Functional territories are based on commuting flows between municipalities, using cluster analysis. These units help avoiding problems common with administrative units, e.g. commuting, as people could travel back and forth without migrating (Cazzuffi and Pereira-López, 2016).

might have offset an *ex-ante* downward trend in migration in Mexico.[18]

	2002			2005		
	Treated	Control	t-test	Treated	Control	t-test
Outcome variables (period end)						
Migrated between waves	.052	.085	-4.62***	.051	.061	-1.40
Internal	.042	.080	-5.78***	.046	.059	-1.87*
International	.011	.005	2.33**	.010	.005	1.92*
Control variables (period start)						
Age	38.68	37.333	4.28***	42.885	42.006	2.55**
Male	.478	.470	0.63	.464	.457	0.41
Years of schooling	5.627	8.288	-24.99***	5.417	8.031	-21.96***
Household head						
Age	47.299	45.726	4.32***	49.534	49.017	1.30
Male	.859	.817	4.23***	.855	.799	4.87***
Years of schooling	4.943	7.399	-22.05***	4.973	7.22	-18.04***
Indigenous	.217	.074	15.01***	.225	.081	13.37***
Household dependency ratio						
0-7	.143	.134	2.02**	.129	.116	3.30***
8-14	.158	.136	4.69***	.144	.129	3.00***
15-20	.114	.099	3.66***	.121	.106	3.27***
66 and more	.037	.030	2.43**	.041	.036	1.65*
Household economic shock	.360	.283	5.98***	.249	.259	-0.71
Household wealth index	-.347	.541	-23.71***	-.258	.580	-19.94***
Lagged marginalization index	-.301	-1.502	66.29***	-.394	-1.683	68.43***
Observations	2,471	2,702		2,094	2,164	

Notes: Means of variables of interest of unbalanced panel estimation sample of 9,431 21-65-year-old individuals. Means of *health status*, *spent on health* and *health expenditures* variables are respectively based on 8,860, 9,183 and 9,099 observations. *** p<0.01, ** p<0.05, * p<0.1.

Table 3 presents descriptive statistics by migration status. On average, migrants are younger, more likely to be men and more educated than non-migrants. They come from households with a greater share of 0-7-year-old dependants, but with a lower share of 8-14-year-old, 15-20-year-old or 66+ dependants. Their households are more likely to have experienced economic shocks in the last 12 months, and to reside in wealthier areas. Disaggregating statistics by treatment status, Table 3 further shows that, on average, migrants in treated areas are older and less educated than migrants in control areas. They come from poorer, more vulnerable families with a slightly higher share of below 15-year-old and 66+ dependants, and are located in more marginalized areas, compared to households with migrants living in control municipalities. These statistics seem to point to affiliation to *Seguro Popular* increasing the probability of those less educated, who might have (more) limited job opportunities, with greater time constraints and coming from poorer, more vulnerable households, to migrate. This is as if publicly provided healthcare enabled families vulnerable to adverse shocks and who could not afford otherwise to send members away by relaxing their financial and care constraints.

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[18] Of the information gathered, migration seems to be largely intra-state and undertaken for job purposes. Unfortunately, data on migration variables could not give consistent information. They are available on request.

Table 3: Descriptive statistics of estimation sample by migration status

| | Migrants | Non-migrants | | Migrants | | | Non-migrants | | |
	Mean	Mean	t-test	Treated Mean	Control Mean	t-test	Treated Mean	Control Mean	t-test
Control variables (period start)									
Age	35.089	40.322	-10.81***	36.547	34.136	2.72***	40.83	39.834	4.08***
Male	.541	.463	3.71***	.564	.526	0.89	.467	.459	0.69
Years of schooling	8.054	6.816	7.21***	6.564	9.028	-7.78***	5.474	8.105	-32.11***
Household head									
Age	44.253	47.975	-6.73***	45.915	43.166	2.33**	48.456	47.512	3.41***
Male	.839	.832	0.48	.856	.828	0.90	.857	.807	6.39***
Years of schooling	6.898	6.127	4.35***	5.581	7.759	-6.39***	4.922	7.284	-27.57***
Indigenous	.099	.150	-3.41***	.165	.055	4.46***	.224	.079	19.47***
Household dependency ratio									
0-7	.170	.128	6.10***	.169	.171	-0.07	.135	.121	4.07***
8-14	.123	.143	-2.90***	.136	.114	1.63	.152	.134	5.09***
15-20	.078	.111	-5.39***	.080	.070	1.86*	.119	.105	4.40***
66 and more	.027	.036	-2.18**	.027	.027	-0.04	.040	.033	2.91***
Household economic shock	.343	.287	2.96***	.386	.316	1.76*	.305	.269	3.80***
Household wealth index	.175	.138	0.61	-.349	.517	-7.67***	-.304	.562	-29.94***
Lagged marginalization index	-1.113	-.974	-3.69***	-.422	-1.564	24.01***	-.340	-1.584	90.92***
Observations	597	8,834		236	361		4,329	4,505	

Notes: Means of variables of interest of unbalanced panel estimation sample of 9,431 21-65-year-old individuals. Means of *health status*, *spent on health* and *health expenditures* variables are respectively based on 8,860, 9,183 and 9,099 observations. *** p<0.01, ** p<0.05, * p<0.1.

5 Results

5.1 Benchmark results

Estimates of equation (1) on an unbalanced panel of 21-65-year-olds are presented in Table 4. Column (1) indicates that the probability of migrating between waves increased by 2.36 percentage points for respondents in municipalities that experienced a significant change in *Seguro Popular* coverage compared to municipalities that did not. This estimate is robust to the inclusion of individual instead of municipality fixed effects, with a slight decrease in estimate magnitude and significance, which could be explained by the reduction in sample size. With municipality fixed effects and controlling for individual and household time-invariant and time-varying variables, from columns (3) to (4), and municipality time-varying variables, in column (5), the estimates similarly give an effect between 2.17 to 2.38 percentage points.

Table 4 clearly points to the expansion of *Seguro Popular* increasing the likelihood to migrate. Looking across all benchmark specifications, the introduction of non-contributory healthcare appears to increase migration by an average of about two percentage points. Compared with a level of 7.4 percentage points in control municipalities, these point estimates suggest an increase in migration of about 30 per cent of the level in control municipalities. In addition, age, whether household heads have indigenous origins and the proportion of dependants in a household decrease the likelihood of a prime-aged member to migrate. Being a man, the occurrence of an economic shock and lagged municipality marginalization index increase the propensity to migrate.

Benchmark results suggest that the expansion of *Seguro Popular* is different from contributory schemes that might be tying affiliates to formal employment, and hence to a specific location. By linking social protection to formal employment, Mexico's contributory social protection system has been shown to increase beneficiaries' security, and thus to reduce the need to migrate, since formal employment increases income stability, which might decrease the necessity to diversify income sources. In contrast, *Seguro Popular* might act as an unconditional cash transfer programme, such as the South Africa Old Age Grant, by which the

reduced occurrence and duration of health shocks, and the alleviation of budget constraints, might free caregivers' time. This could enable them to reallocate across space and diversify income sources, while ensuring coverage of household dependants. The relatively small magnitude of these coefficient estimates might be expected because access to non-contributory healthcare does not induce a stable source of income, but rather punctual transfers to compensate out-of-pocket health expenditures. In this sense, the effect of publicly provided healthcare might be close to the impact of small monetary incentives on internal migration in rural Bangladesh, as found in Bryan et al. (2014).

Table 4: Coefficient estimates of benchmark specifications

Variables	(1)	(2)	Has migrated (3)	(4)	(5)
Treated X 2005	0.0236**	0.0176*	0.0228**	0.0238**	0.0217**
	(0.0098)	(0.0104)	(0.0098)	(0.0098)	(0.0098)
2005	-0.0234***	-0.0112	-0.0134*	-0.0123*	-0.0064
	(0.0074)	(0.0081)	(0.0074)	(0.0074)	(0.0078)
Treated		-0.0112			
		(0.7075)			
Age			-0.0021***	-0.0018***	-0.0018***
			(0.0003)	(0.0003)	(0.0003)
Male			0.0167***	0.0147***	0.0146***
			(0.0051)	(0.0052)	(0.0052)
Years of schooling			0.0011	0.0015	0.0015
			(0.0008)	(0.0011)	(0.0011)
Age of head				-0.0005	-0.0005
				(0.0003)	(0.0003)
Head is male				-0.0009	-0.0008
				(0.0070)	(0.0070)
Years of schooling of head				-0.0008	-0.0007
				(0.0010)	(0.0010)
Head is indigenous				-0.0234**	-0.0237**
				(0.0101)	(0.0101)
0-7 dependency ratio				-0.0030	-0.0033
				(0.0211)	(0.0211)
8-14 dependency ratio				-0.0729***	-0.0727***
				(0.0169)	(0.0169)
15-20 dependency ratio				-0.0648***	-0.0645***
				(0.0171)	(0.0171)
66 and more dependency ratio				-0.0488*	-0.0483*
				(0.0273)	(0.0272)
Household economic shock				0.0218***	0.0220***
				(0.0059)	(0.0059)
Household wealth index				-0.0032	-0.0034
				(0.0022)	(0.0022)
Lagged marginalization index					0.0321**
					(0.0133)
Municipality FE	Yes	No	Yes	Yes	Yes
Individual FE	No	Yes	No	No	No
Mean of dependent variable	.0633	.0566	.0633	.0633	.0633
	(.2435)	(.2311)	(.2435)	(.2435)	(.2435)
Observations	9,431	7,118	9,431	9,431	9,431
R-squared	0.0226	0.5598	0.0343	0.0401	0.0404

Notes: Estimates are for 21-65-year-old individuals, interviewed in at least two consecutives waves. The dependent variable is a binary variable taking unity if an individual migrated between wave 1 (2002) and wave 2 (2005) and/or between wave 2 and wave 3 (2009). Columns (1)-(5) present coefficient estimates of linear probability models. Standard errors robust to heteroscedasticity are reported in parentheses. *** $p<0.01$, ** $p<0.05$, * $p<0.1$.

5.2 Robustness checks

Several tests are conducted to assess the validity of the identifying assumptions. An important threat to identification would be a significant relationship between municipality-specific timing of *Seguro Popular* roll-out and migration trends – differential time trends between treated and control municipalities correlated with its expansion. This would bias the estimated average treatment effects. Because there are only two time periods (2002-2005 and 2005-2009), municipality-specific trends cannot be controlled for. In addition to controlling for time-varying characteristics at the municipality level by including lagged marginalization indices, placebo tests are performed. As in Alcaraz et al. (2016) and de Janvry et al. (2015), a first placebo test is run by estimating equation (1) on first-period data with treatment status as variable of interest. The change in migration between wave 1 and wave 2 is regressed on treatment status, controlling for individual, household and municipality characteristics, as in:

$$Y_{im02/05} = \alpha + \beta_1 Treated_{m05} + \delta' X_{imt} + u_{imt} \tag{2}$$

Column (1) of Table 5 indicates that the propensity to migrate was significantly but negatively related with a major expansion of *Seguro Popular* in the next period. Although migration propensities between treated and control municipalities before programme expansion were different, the fact that the sign on treatment status is opposite to difference-in-differences estimates of Table 5 supports the identification strategy.

A second placebo test is performed by randomly assigning the treatment to control municipalities. Specifically, (i) the estimation sample is restricted to control municipalities; (ii) a 'fake' treatment is randomly assigned to half of the control municipalities; (iii) equation (1) is estimated; and (iv) this procedure is repeated (bootstrapped) 1,000 times. As shown in column (2) of Table 5, the difference-in-differences estimator is statistically significant and negative, suggesting that the difference-in-differences estimates of benchmark specifications do not reflect the existence of any positive selection into the treatment and/or pre-programme positive trend in treated municipalities. Its negative sign indicates that benchmark difference-in-differences estimates represent the lower bound of the true effect of the expansion of *Seguro Popular* on migration.

Another threat to identification would be that the timing of *Seguro Popular* expansion is associated with significant changes in the probability to migrate before its introduction. For instance, if *Seguro Popular* was expanded to react to (pre-programme) downward trends in migration, estimates could mirror what was intended, i.e. changes to average migration rates. Individuals and households could also have anticipated that they would benefit from a greater coverage and lowered their propensity to migrate before its expansion. In this case, estimated effects would reflect returns to 'normal' migration rates. Since there are only two time periods, the robustness of the estimates to a potential pre-treatment 'trend' specific to treated observations – an Ashenfelter dip effect (Ashenfelter, 1978) – cannot be assessed. However, it is reasonable to assume that increasing migration rates have not driven the expansion of *Seguro Popular*, as policies tend to fight rather than encourage migration, in particular internal migration, fearing unwieldy, unsustainable urbanisation. Moreover, focusing on municipalities that experienced changes in coverage in its expansion phase, not in its pilot phase, is assumed to rule out the existence of households' or individuals' anticipatory migration behaviours.

A last threat to identification is the attrition of households and individuals from the MxFLS. Around 51% of estimation sample observations were not successfully interviewed in all three waves. Estimates would be biased if there were selection into attrition (retention) due to the expansion of *Seguro Popular*. The probability of estimation sample respondents not to be interviewed in all three waves of the survey is first regressed. As column (3) of Table 5 shows, the effect of a change in coverage is statistically significant and negative. This is evidence of negative (positive) selection into attrition (or retention). Those living in municipalities that experienced a significant change in coverage are more likely to be successfully interviewed in all three waves. Equation (1) is then run on a balanced panel. The difference-in-differences estimate in column (4) of Table 5 is of a similar magnitude, but slightly loses in statistical significance compared

to difference-in-differences estimates with the unbalanced panel. This suggests that, despite potential selection, panel attrition might not substantially affect the estimated effect of *Seguro Popular* on respondents' propensity to migrate.

<div align="center">Table 5: Robustness checks</div>

| Variables	Has migrated (1)	Has migrated (2)	Attrition (3)	Has migrated (4)
Treated X 2005			-0.0321**	0.0267*
			(0.0149)	(0.0155)
Fake treatment X 2005		-0.0612***		
		(0.0006)		
2005		0.0916***	-0.0774***	-0.0139
		(0.0007)	(0.0110)	(0.0136)
Treated	-0.0343***			
	(0.0094)			
Time FE	No	Yes	Yes	Yes
Municipality FE	No	Yes	Yes	Yes
Control variables FE	Yes	Yes	Yes	Yes
Mean of dependent variable	.0692	.0742	.5106	.0594
	(.2538)	(.0001)	(.4999)	(.2363)
Observations	5,173	4,866	9,431	4,616
R-squared	0.0263	0.0393	0.4819	0.0590

Notes: Estimates are for 21-65-year-old individuals, interviewed in at least two consecutive waves in columns (1)-(3); in three consecutives waves in column (4) (balanced panel). In column (1), the estimation sample is limited to the first time period. In column (2), observations are limited to never treated that were assigned a fake treatment. In columns (1), (2) and (4), the dependent variable is a binary variable taking unity if an individual migrated between wave 1 (2002) and wave 2 (2005) and/or between wave 2 and wave 3 (2009). In column (3), the dependent variable is a binary variable that takes value 1 if a respondent with non-missing information was not successfully interviewed in three consecutive waves. Columns (1)-(4) present coefficient estimates of linear probability models. Column (1) presents estimates of a placebo test; column (2), of a falsification test; columns (3)-(4) investigates panel attrition. Standard errors robust to heteroscedasticity are in parentheses. In column (2), random assignment of fake treatment and regressions were bootstrapped (1,000 repetitions) (standard errors reported in parentheses in column (2) are bootstrapped). *** p<0.01, ** p<0.05, * p<0.1.

5.3   Heterogeneous effects

The probability to migrate internally and internationally are separately regressed on the full estimation sample. As columns (1) and (2) of Table 6 show, difference-in-differences estimates only hold for internal migration, with a statistically significant increase of 2.42 percentage points. The effect on international migration is null but statistically insignificant. The fact that the insignificance of the expansion of *Seguro Popular* on international migration is explained by the very low average international migration in the estimation sample (0.75%) cannot be ruled out. However, it can be the case that access to healthcare has a significant effect on internal migration but insignificant on international migration, in particular if difficult access to financial capital and budget constraints have been limiting migration. This might be because internal migration tends to be less expensive, less risky than international migration, and because affiliation to *Seguro Popular* does not directly provide cash, but increases disposable income by limiting health expenditures.

The full estimation sample is then decomposed by gender to account for some degree of gender-differentiated time and task distribution. In Mexico, women tend to spend more time caring for dependants than men (del Valle, 2016). Columns (3) and (4) of Table 6 confirm this hypothesis: men are significantly more likely

to migrate than women following a change in coverage. This is consistent with evidence from South Africa and India showing that, when women are those affiliated, other family members, in particular men, tend to migrate.[19] Treated households might simultaneously follow different livelihood strategies. Some household members, women, would stay home to benefit from local labour market opportunities and affiliation to *Seguro Popular*, while taking care of dependants when they do not work outside their households. Men, now financially 'enabled' to leave, with less time tied to dependants and not socially expected to care for them, would migrate (Hagen-Zanker and Leon-Himmelstine, 2013).

Table 6: Heterogeneous effects

	Has migrated			
	Internal	International	Female	Male
Variables	(1)	(2)	(3)	(4)
Treated X 2005	0.0242**	0.0000	0.0152	0.0283*
	(0.0094)	(0.0035)	(0.0125)	(0.0156)
2005	-0.0031	-0.0008	-0.0073	-0.0037
	(0.0076)	(0.0023)	(0.0098)	(0.0124)
Time FE	Yes	Yes	Yes	Yes
Municipality FE	Yes	Yes	Yes	Yes
Control variables	Yes	Yes	Yes	Yes
Mean of dependent variable	.0578	.0075	.0546	.0732
	(.2334)	(.0864)	(.2272)	(.2605)
Observations	9,431	9,431	5,017	4,411
R-squared	0.0404	0.0421	0.0488	0.0410

Notes: Estimates are for 21-65-year-old individuals, interviewed in at least two consecutive waves. In columns (1)-(4), the dependent variable is a binary variable taking unity if an individual migrated between wave 1 (2002) and wave 2 (2005) and/or between wave 2 and wave 3 (2009). In column (1), the dependent variable takes value 1 if migration was internal; 0, otherwise. In column (2), it takes value 1 if migration was international; 0, otherwise. In columns (3), the estimation sample is limited to women; in column (4), to men. Columns (1)-(4) present coefficient estimates of linear probability models. Standard errors robust to heteroscedasticity are in parentheses. *** p<0.01, ** p<0.05, * p<0.1.

Beneficiaries' characteristics matter in explaining these dynamics. Following Angelucci (2015), Figure 5 depicts pre-programme distributions of years of schooling (left), dependency ratios (right) and household wealth (bottom) for non-migrants in control municipalities, migrants in control municipalities and migrants in treated municipalities. The migrant skill distribution in control municipalities has more density in its middle (dashed grey line) than non-migrants (solid grey line), but it is shifted to the right compared to that of migrants in treated municipalities. As in the case of *Oportunidades* and international migration in Angelucci (2015) or in Greenwood and McDowell (2011), the figure on the left indicates that migrants are negatively selected (into migration) with regard to education. By alleviating financial constraints on those who are the most likely to be affected by health shocks, significant changes in health insurance coverage might worsen migrants' skill profiles, since, if skill set is a proxy for labour market opportunities, unskilled migrants are those facing the greatest difficulties in funding migration. Accessing health insurance might thus enable those with a relatively bad skill set to expand their work opportunities across space. While the figure at the bottom suggests that migrants in treated municipalities are poorer than migrants in control

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[19] See for instance Sienaert (2008) for South Africa.

municipalities, the statistical relationship between either dependency ratios or wealth is not confirmed by Mann-Withney tests.[20]

Figure 5: Years of schooling, dependency ratio and wealth index by treatment status
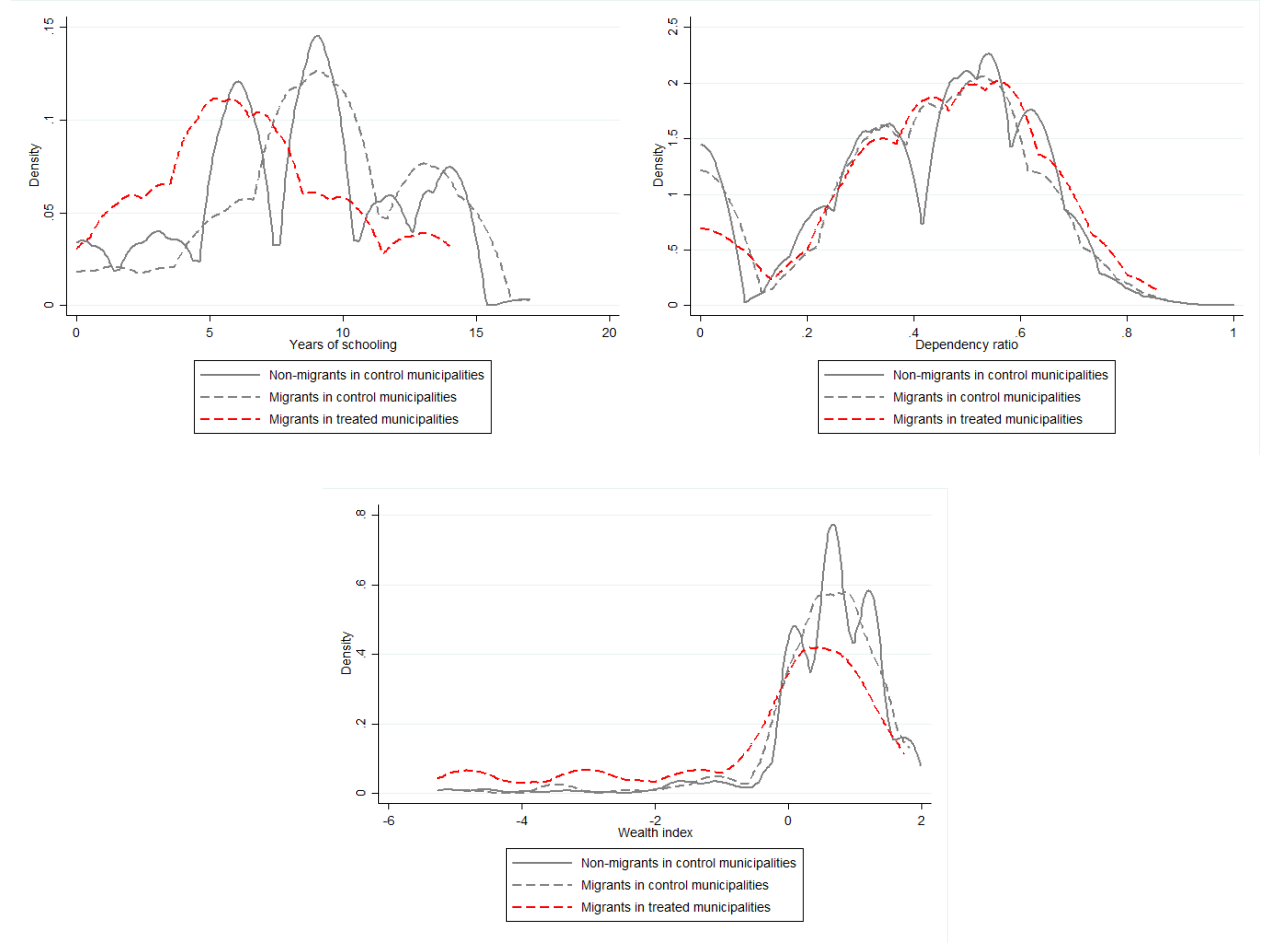
5.4 Potential mechanisms

This subsection further sheds light on mechanisms that might be at stake by running equation (1) on a different set of outcomes. Column (1) of Table 7 suggests that a change in healthcare coverage has an almost null, insignificant effect on subjective health reported at the end of each time period. Columns (2) and (3) look at the relationship between change in coverage and time dedicated to caring for dependants. Column (2) shows that a major increase in publicly-provided healthcare decreases the propensity to spend time caring. The coefficient estimate on the number of hours in column (3) is negative but statistically insignificant. Column (4) indicates that changes in coverage are associated with a greater propensity to experience, or to report, health-related economic shocks in a household. This is in line with the fact that, to benefit from non-contributory healthcare, health-related shocks have to be reported. Estimates from columns (1)-(4) taken together suggest that healthcare expansion has improved households' resilience to health-related shocks, and enabled caregivers to reallocate time from caring to work outside their household, thus expanding their work opportunities across space.

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[20] Estimates are available on request.

The increase in disposable income following affiliation to *Seguro Popular* could not only limit financial constraints, but also, if used as collateral, increase beneficiaries' credit worthiness. Loans could then fund migration. Equation (1) is run with a dependent variable that takes value 1 if respondents have asked for a loan in the 12 months preceding the end of the time period of each wave.[21] Although this is an imperfect measure of borrowing behaviours,[22] column (5) of Table 7 shows a positive but statistically insignificant relationship between changes in healthcare coverage and likelihood to borrow in the previous 12 months. While the fact that this insignificance is due to the relatively high variance of the outcome variable cannot be ruled out, it does not clearly support the hypothesis.

In columns (6) and (7) of Table 7, equation (1) is run with a dependent variable that takes value 1 if respondents' households have made health expenditures in the three months preceding the end of each time period (6), and how much they spent on health (7). Column (6) suggests a positive and statistically significant relationship between changes in healthcare coverage and likelihood to spend on health; column (7), a positive but insignificant association between changes in coverage and amount spent on health expenditures. While estimates of column (7) pass robustness checks, those of column (6) do not, suggesting that individuals living in both treated and control areas were more likely to spend on health expenditures in the second time period compared to the first (likelihood). However, respondent living in control municipalities tended to spend comparatively less money on health (quantity).[23]

Taken together, estimates suggest that changes in *Seguro Popular* affiliation did not affect migration by relaxing financial constraints through borrowing or health expenditures, but rather by freeing up working-age members in affiliated households from care constraints. Induced increases in disposable income might not be substantial enough, and/or internal migration might be cheap enough not to require a lot of financial means. In this setting, time rather than financial constraints might be binding domestic migration.

Table 7: Investigating mechanisms

Variables	Health status (1)	Spent time caring (2)	Hours spent caring (3)	Health-related shock (4)	Borrowed money (5)	Spent on health (6)	Health expenditures (7)
Treated X 2005	.0037	-.0408**	-.3295	.0331**	0.0129	.0072*	5.4260
	(.0100)	(.0161)	(.5997)	(.0136)	(0.0125)	(.0039)	(71.7706)
2005	-.0034	.0655***	1.2903***	.0219**	0.0547***	-.0017	38.0591
	(.0068)	(.0123)	(.4546)	(.0105)	(0.0097)	(.0034)	(43.7340)
Time FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Municipality FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Control variables	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Mean of dependent variable	.9470	.2231	5.8334	.1225	.1003	.9909	349.0111
	(.2241)	(.4163)	(15.3044)	(.3278)	(.3004)	(.0952)	(1701.176)
Observations	8,860	9,431	9,431	9,431	9,431	9,183	9,099
R-squared	0.0396	0.1578	0.1402	0.0382	0.0464	0.0183	0.0361

Notes: Estimates are for 21-65-year-old individuals, interviewed in at least two consecutive waves. In column (1), the dependent variable is a binary variable taking unity if an individual reported to have regular, good or very good health at the end of the time period; 0, if s/he reported bad or very bad health. In column (2), the dependent variable is a binary variable taking unity if an individual dedicated time to caring for dependants over the previous week at the end of the time period. In column (3), the dependent variable is a continuous variable measuring the number of hours an individual dedicated to caring for dependants over the previous week at the end of the time period. In column (4), the dependent variable is a binary variable taking value 1 if an individual belongs to a household that experienced at least one health-related economic shock in the five years preceding the end of the time period. In column (5), the dependent variable is a binary variable that takes value 1 if an individual has borrowed in the 12 months preceding the end of the time period. In column (6), the dependent variable is a binary variable that takes value 1 if the household of respondents has had health expenditures in the 3 months preceding the end of the time period. In column (7), the dependent variable is a continuous variable measuring how much a household of respondents has spent on health expenditures in the 3 months preceding the end of the time period. Columns (1)-(5) present coefficient estimates of linear probability models. Standard errors robust to heteroscedasticity are in parentheses. *** p<0.01, ** p<0.05, * p<0.1.

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[21] That is up to a year from 2005 for the first time period, and up to a year from 2009 for the second time period.

[22] A household member might ask for a loan for another relative. Loans that were used to fund migration that took place more than 12 months before MxFLS 2 and MxFLS 3 are not reported.

[23] Estimates are available on request.

6 Conclusion

While the academic and political debate has emphasised adverse labour market incentives, the positive effects of safety nets on health and poverty are increasingly recognised. A growing literature has been studying how social protection programmes relate to household livelihood strategies, in particular the decision to migrate within recipient households. A contribution of this paper is to exploit the expansion of a publicly provided healthcare programme initiated in 2002 in Mexico to obtain causal estimates on this relationship.

This paper exploits the timing and the panel structure of the MxFLS and uses a difference-in-differences specification to estimate a non-negligible migration effect in municipalities that experienced a major change in coverage rate in the middle of *Seguro Popular* roll-out. Individuals living in treated municipalities were more likely to migrate following a change in exposure to the treatment, compared to respondents living in municipalities that did not experience a significant change in coverage rate. Examining estimate heterogeneity suggests that associated increases in disposable income were not substantial enough to fund international migration, in contrast to other non-contributory social protection schemes, e.g. conditional cash transfer programmes. The increase in the probability to migrate is only statistically significant for men, supporting the idea that, in a context of gender-differentiated task distribution and income source diversification like in Mexico, men are those more likely to migrate compared to women. Robustness checks confirm the validity of the identification strategy against threats of time-trending unobservables that might vary significantly between treated and control municipalities. They reveal that changes in migration propensity prior to the programme were negatively correlated with its expansion, consistent with current migration trends within and from Mexico.

In showing that (non-contributory) safety nets can increase the propensity to migrate, these results shed light on some (unattended) effects of publicly-provided healthcare. They suggest that migrating might be a channel through which labour market behaviours and livelihood strategies are affected. Building financial strength and freeing up caregivers' time through reduced occurrence and duration of health shocks can encourage labour force detachment of working-age members in households vulnerable to adverse shocks. By enabling the spatial reallocation of labour available within a household, having coverage furthers the diversification of household income sources, which is likely to help families breaking out of poverty traps.

While findings are obtained for the case of Mexico, with about 9,000 observations from household survey data that are not representative at the municipality level (treatment unit), which might threaten the external validity of this study, they have interesting implications. Given the importance publicly provided healthcare has been receiving as a means to reduce poverty, while potentially distorting labour markets, analysing dynamics between access to healthcare and migration is likely to be at the centre of the social policy debate. This paper contributes to this discussion by providing causal estimates of the impacts of a significant change in publicly provided healthcare on the propensity to migrate. It also suggests that it is necessary to include both recipients and household members living with them *or* who have migrated in this analysis. Results emphasise the importance of including household members who do not reside with recipients. Not accounting for the effects of social protection programmes on migration might question the reliability of results obtained in other literatures, such as labour market behaviours.

This paper only focused on short-run effects. However, healthcare coverage has been found to have positive effects on the health and educational outcomes of dependants, in particular children of affiliates (Alcaraz et al., 2016). For this reason, these dependants could be expected to display a greater propensity to migrate in the longer run if, while they are equipped with better health and higher education, local labour markets did not offer them adequate opportunities. Investigating the implications of non-contributory healthcare expansion on the likelihood to migrate in the long term is thus an interesting direction for research.

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