Economic Inquiry

IMMIGRANT NETWORKS AND THE TAKE-UP OF DISABILITY PROGRAMS: EVIDENCE FROM THE UNITED STATES

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We examine the role of ethnic networks in disability program take-up among working-age immigrants in the United States. Controlling for country of origin and area of residence fixed effects, immigrants residing amid a large number of co-ethnics are more likely to receive disability payments when their ethnic groups have higher take-up rates. Differences in satisfying the work history or income and asset requirements of the disability programs explain part of this relationship, but social norms also play an important role. Information sharing appears influential for Supplemental Security Income take-up but not for Social Security Disability Income. (*JEL* J61, H55, I18)

I. INTRODUCTION

In 2009, the two largest disability programs in the United States, namely the Social Security Disability Insurance (DI) program and the Supplemental Security Income (SSI) disability program, paid approximately 158.9 billion dollars in benefits to the disabled (U.S. Census Bureau 2012, Tables 547 and 563).[1] Interestingly, despite improvements in the overall health of the population over time, the two programs have grown substantially both in terms of benefits per recipient and number of recipients (Autor and Duggan 2006; Social Security Administration 2006). A recent Congressional Budget Office (CBO) report projects that the DI trust fund will be exhausted by 2017 if no legislative actions are taken (CBO 2014). As policy-makers evaluate potential changes to these programs, important considerations include whether benefits are currently being awarded optimally and how any policy changes may ultimately impact disability program take-up. To gain insight into these issues, this paper explores how networks, specifically ethnic networks, affect the probability that immigrants receive disability payments either from DI or SSI.

If eligibility for the disability programs were exogenously determined, Social Security examiners were perfectly able to distinguish between who is and who is not able to work, and everyone who was eligible for the programs applied for and ultimately received benefits, then we would not expect social networks to play a

*For the many valuable comments and suggestions, we would like to thank each of our referees as well as seminar participants at Boston University, Institute for the Study of Labor (IZA), Lancaster University, Louisiana State University, Northeastern University, University of Cyprus, University of Connecticut, Vassar College, UMass-Amherst, and University of Colorado at Denver and conference participants at the Yale University Health Workshop, EALE in Cyprus, PAA meetings in San Francisco, NARSC meetings in Miami, and the SOLE meetings in Boston. We are especially indebted to Madeline Zavodny and Pia Orrenius for providing the occupational injury data. This project was started while Delia Furtado was visiting the Yale School of Public Health on an Agency for Health Care Research and Quality (AHRQ) Fellowship. The research reported here was performed pursuant to a grant from the US Social Security Administration (SSA) funded as part of the Retirement Research Consortium (RRC). The opinions and conclusions expressed are solely those of the authors and do not represent the opinions or policy of SSA, any agency of the federal government, the RRC, the University of Connecticut, the University of Cyprus, or Boston College.

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1. In comparison, only about 10.5 billion dollars were paid to Temporary Assistance for Needy Families (TANF) recipients in the same year (U.S. Census Bureau 2012, Table 567).

ABBREVIATIONS

ACS: American Community Surveys
BLS: Bureau of Labor Statistics
CBO: Congressional Budget Office
DDS: Disability Determination Services
DI: Social Security Disability Insurance
IPUMS: Integrated Public Use Microdata Series
NHIS: National Health Interview Survey
PUMAs: Public Use Microdata Areas
SSA: Social Security Administration
SSI: Supplementary Security Income
WVS: World Values Survey

Economic Inquiry
(ISSN 0095-2583)
Vol. 54, No. 1, January 2016, 247–267

doi:10.1111/ecin.12248
Online Early publication August 6, 2015

strong role in disability program take-up. On the other hand, if the Social Security Administration (SSA) was not able to screen applicants perfectly, then among those with marginal disabilities, ultimate decisions about applying for benefits may depend on social norms regarding exaggerating disabilities or the benefits of leisure. Even if the SSA was able to perfectly screen applicants, stigma against receiving government benefits while not working may prevent the genuinely disabled from claiming benefits. Moreover, if the application process is sufficiently complex, then information sharing within social networks may be an important determinant of take-up among the truly disabled.[2] Regardless of exactly how networks operate, their existence implies that any policy which would change the number of people eligible for benefits might have substantial multiplier effects.

Network effects are notoriously difficult to estimate empirically (Manski 1993). We can show that individual disability program take-up is positively correlated with average disability program take-up in a person's neighborhood, but this may simply reflect unobservables which vary by neighborhood. Another approach in identifying networks might be to examine the relationship between individual disability program take-up of immigrants and average take-up by country of origin, but this also cannot be taken as proof of networks as there might be differences in the tendency to become disabled which vary by country of origin.

To address these issues, we use an empirical approach similar to the one pioneered by Bertrand, Luttmer, and Mullainathan (2000) in their study of welfare take-up. They find that being surrounded by people who speak the same language increases welfare use more for people in high welfare-receiving language groups, a result they interpret as evidence of network effects. Aizer and Currie (2004) take a comparable approach to identify the role of networks in the use of publicly funded prenatal care. Other researchers have found evidence of network effects in health care utilization (Deri 2005; Devillanova 2008), Medicaid take-up (Gee and Giuntella 2011), and WIC participation during pregnancy (Figlio, Hammersma, and Roth 2011). Åslund and Fredriksson (2009) estimate the same equation as Bertrand et al. but exploit the plausibly exogenous placement of refugees in Sweden for tighter identification.

Much of the literature on network effects examines participation in transfer programs aimed at the poor. In our analysis, we start by examining DI, an insurance program requiring awardees to pay into the system for several years before becoming eligible for payments in the event of a work-preventing disability. Taboos against take-up of this type of program may be less strong than one aimed at people who have spent years out of the labor market. For purposes of comparison, we also study network effects in the take-up of disability-related SSI, a program without the work history requirements of DI but with income and asset limits.

Rege, Telle, and Votruba (2012) are among the first of the analyses of the role of social interactions in disability program participation. Using neighbors' exposure to plant downsizing as an instrument for disability program participation, the authors find that Norwegians living geographically close to people who participate in the program are more likely to receive disability payments themselves. Exploiting random assignment of judges to DI applicants whose cases are initially denied, Dahl, Kostøl, and Mogstad (2014) find that when a parent is allowed DI at the appeal stage, there is an increased likelihood that their adult children participate in the program. Not only does our paper differ from these papers in terms of empirical approach, but also in our focus as it is on immigrant networks within a U.S. context. We also examine two types of disability programs aimed at different populations.

In a companion paper (Furtado and Theodoropoulos 2013), we use National Health Interview Survey (NHIS) data to explore the role of ethnic networks on the take-up of SSI. We find that immigrants who are more exposed to co-ethnics receiving SSI for a disability are themselves more likely to apply for the program, but conditional on applying, they are also more likely to ultimately be denied benefits. This implies that when immigrants are exposed to more SSI take-up within their communities, they are more likely to apply with marginal disabilities.

Our analysis of U.S. Census 2000 data provides evidence of social interaction effects for both DI and SSI take-up. Immigrants living in neighborhoods with many others from the same origin country are especially likely to receive DI

2. Social norms may work in conjunction with information sharing if network members share information about doctors who are most likely to exaggerate disabilities. According to a *New York Times* article, three doctors were responsible for 86% of Long Island Railroad's disability applications. They were charged with preparing fraudulent medical assessments for hundreds of retirees (Raushbaum and Secret 2011).

benefits if they belong to high DI ethnic groups. The relationship is even stronger for SSI. Results are robust to adding a series of controls for assimilation, human capital, and disability to the model suggesting that the country of origin and area of residence fixed effects are effectively controlling for the most egregious sources of bias. We also construct for each country of origin-local area cell, unemployment rates, average wages, on-the-job injury rates, on-the-job fatality rates, average age, average years of schooling, and average years in the United States. Adding these variables to the model has no impact on our estimated network effects.

We then explore how ethnic networks operate. To measure home country norms, we use a variety of questions from the World Values Survey (WVS). While the country of origin fixed effects will incorporate direct impacts of home country norms on disability program take-up, we show that immigrants from countries with strong government cheating taboos and importance of work norms are less responsive to exposure to DI and SSI take-up within their ethnic communities. This certainly points to a potential role of social norms in explaining our estimated network effects. Further, we show that information sharing might be playing a role for SSI take-up but not DI take-up and that leisure complementarities are not driving network effects in either program.

A remaining issue when interpreting these findings, however, is that immigrants residing amid a large number of co-ethnics may have unobservable characteristics which more closely resemble the average characteristics of group members. Of particular concern is whether immigrants from the same country living near each other have similar likelihoods of meeting the nondisability requirements for DI or SSI. Recall that in addition to a work-preventing disability, applicants must have sufficient work histories to qualify for DI and meet certain income and asset limits to qualify for SSI. If immigrants with extensive work histories are more likely to live in ethnic enclaves when they belong to groups that tend to have long work histories, then we might observe correlations in their DI take-up not because they are more likely to apply for the program (conditional on qualifying) but simply because they are more likely to qualify. Similarly, if immigrants living below the poverty line are more likely to live in ethnic enclaves when they belong to high poverty ethnic groups, correlations in SSI take-up might be driven by differences in poverty rates as opposed to norms or information sharing regarding the SSI program per se.

To examine how problematic this is, we exploit the fact that *regardless of disability*, people aged 65 and above are eligible for Social Security retirement income as long as they satisfy the program's work history requirements and are eligible for SSI if they satisfy the income and asset requirements. Given that social norms about exaggerating a disability and information sharing about the appeals process do not play any role in the decisions to apply for retirement income for these older immigrants, we interpret positive and statistically significant network coefficients in this population as evidence that part of our estimated network effects in the baseline models are driven by differences in satisfying the nondisability-related requirements of the programs.

We find statistically significant but substantially smaller estimated network effects in our retirement-age sample suggesting that while eligibility differences are important, they are not the sole drivers of our results. In addition, we find no evidence that home country social norms measured in the WVS have an impact on this population. Taken together, these results suggest that social norms do affect how exposure to disability program participation within ethnic communities translates into disability program take-up in the baseline sample.

The remainder of the paper is organized in the following way. Section II provides background information on the DI and SSI disability programs. Section III explains our identification strategy. Section IV presents the data and Section V outlines the main results and addresses concerns about omitted variable bias. Sections VI and VII examine the mechanisms through which networks operate and conclusions are provided in Section VIII.

II. BACKGROUND ON DISABILITY PROGRAMS IN THE UNITED STATES

The DI program was established in 1956 to insure U.S. workers against the risk of being unable to work due to a physical or mental disability. To be eligible, applicants must satisfy both a "recent work" requirement, which usually amounts to working 5 of the past 10 years for workers over the age of 30 and a "duration of work" requirement, which generally entails working one quarter of the years since turning 21. The SSI program, enacted in 1974, also provides cash benefits to working-age disabled or blind individuals. Although it generally does not have work history requirements, the SSI program does have asset and income limits which vary by state. Thus, while both programs provide cash benefits to the disabled, DI is an insurance program while SSI is a welfare program. A disabled person may receive benefits from both DI and SSI if he or she satisfies the work history requirements of DI, but DI payments are not sufficient to bring the person above the SSI income limits.

The same process is used to determine whether a person is disabled for both programs. First, examiners verify whether or not the individual has engaged in substantial gainful activity, defined in the year 2010 as earning $1,000 per month, in the previous 5 months. Next, they examine the medical evidence to determine whether the impairment is severe enough to prevent work for at least a year or result in death. If the answer is yes, and the condition is on the list of impairments, then benefits are awarded. Applicants with severe disabilities that are not on the list of impairments are also awarded benefits if examiners determine that they are not able to perform any job in the national economy given their age, skills, and work experience. Even when benefits are ultimately denied, there is an extensive appeals process which is often successful.[3] Roughly one-third of all DI applications are awarded initially and about two-thirds of all initial applications are awarded by the end of the appeals process (Maestas, Mullen, and Strand 2013). SSI applications have lower approval rates than DI applications (Annual Statistical Report on the DI Program 2010; SSI Annual Statistical Report 2010).

The DI and SSI programs differ with respect to benefits. DI payments are a function of past earnings. High earners receive more than low earners, but the benefit formula is progressive in that replacement rates are higher for low earners than that of high earners. DI recipients are also eligible for Medicare coverage after 2 years of receiving DI payments. SSI payments are on average lower than DI payments and tend to vary by state of residence because of the way different states supplement federal benefits. SSI recipients are eligible for Medicaid immediately upon being awarded benefits.

Before 1996, legal immigrants were eligible for both DI and SSI as long as they satisfied the other requirements of the programs. The Personal Responsibility and Work Opportunity Reconciliation Act (PRWORA) of 1996 imposed many additional restrictions with respect to SSI eligibility on all noncitizens, including those legally in the United States. Initially, practically all noncitizens were barred from receiving SSI, but later reforms restored SSI disability benefits to those who were legally residing in the United States on August 22, 1996. All immigrants in our sample were residing in the United States 5 years prior to the 2000 Census, and so, as long as they satisfy the other program requirements and are legally residing in the United States, they are eligible for both types of disability programs.[4]

III. EMPIRICAL APPROACH

Estimating social interaction effects is difficult both because information on people's social contacts is typically unavailable and friendships are not formed randomly. It turns out, however, that by making certain assumptions about who is likely to be in people's social circles, we can control for many of the unobserved variables which make it difficult to study network effects.

One often-made assumption in the social interactions literature is that people are more likely to befriend those who live geographically close to them. A researcher might examine whether people who reside amid many others who receive DI or SSI payments are themselves more likely to receive these payments. The problem with this approach is that even in a world with no social interaction between neighbors, a within-neighborhood correlation in disability program participation could result from similar tendencies to become disabled or similarities in labor market opportunities. From a purely bureaucratic perspective, people apply for benefits at their local disability determination services (DDS) offices and so regional variation in the leniency of DDS

3. Rejected applicants can usually ask for reconsideration at the same DDS office. The next level is a hearing before an SSA administrative law judge where the claimant appears in person. Further appeals can be made to the Appeals Council and the federal courts. For a detailed discussion and a graphical representation of the application and appeal process, see Benitez-Silva et al. (1999).

4. Immigrants arriving in the United States after August 22, 1996 can receive SSI benefits if they have strong military connections, long work histories, or are cross-border Native Americans. Refugees and other immigrants admitted for humanitarian reasons are only eligible during their first 7 years in the United States. Other non-citizens cannot receive SSI.

TABLE 1
Percentage of Immigrants Receiving DI or SSI by Country of Origin

DI			SSI		
Top 5	Percentage	Observations	*Top 5*	Percentage	Observations
Cape Verde	4.60	618	Republic of Georgia	9.33	121
Yemen, Arab Republic	4.29	339	Cambodia	7.45	4,043
Croatia	3.99	862	Laos	6.19	5,699
Portugal	3.66	6,717	Belarus	4.96	550
Italy	3.56	9,959	Kosovo	4.81	96
Bottom 5			*Bottom 5*		
St. Kitts-Nevis	0.43	287	Liberia	0.17	660
Sri Lanka (Ceylon)	0.40	595	Northern Ireland	0	271
Northern Ireland	0.38	271	Nepal	0	145
Nepal	0	145	Algeria	0	201
Tanzania	0	247	Tanzania	0	247

Notes: Our sample consists of non-widowed, non-institutionalized immigrants, age 25–61, who are not currently in school and who were living in the United States 5 years prior to the survey. Only naturalized citizens and non-citizens are considered. We also drop American Indians, Alaska natives, and Hawaiians as well as people whose countries of origin are not specified in the data. Percentages are weighted using the appropriate person-level weights provided by the 2000 U.S. Census.

offices could drive the correlation in disability program participation.[5]

An alternative way to proxy for social circles, at least for immigrants, is with country of origin. Immigrants typically arrive to the United States with little knowledge of U.S. customs, institutions, and language, making it easier for them to interact with others from the same country of origin as opposed to natives or immigrants from different countries. While only 1.7% of the 25- to 61-year-old immigrants in our sample receive DI payments, the proportion ranges from 4.6 among Cape Verdeans to 0 among immigrants from Tanzania. The ethnic variation in the proportion receiving SSI is even greater, ranging from 9.3 for people from the Republic of Georgia to 0 for people from Tanzania (see Table 1). Again, however, immigrants from the same country are likely to have similar tendencies to become disabled and may face similar labor market opportunities.

To address these issues, we use an approach pioneered in Bertrand et al.'s study of welfare cultures. Specifically, we assume immigrants are likely to interact predominantly with people from their country of origin who also live within close geographic proximity. We then examine whether immigrants residing amid a large number of co-ethnics are more likely to receive disability payments when their ethnic groups have stronger disability program usage tendencies. We estimate the following equation using a linear probability model:

$$(1) \qquad D_{ijk} = \beta_1 \overline{D}_j \times \mathrm{CA}_{jk} + \beta_2 \mathrm{CA}_{jk} + \mathbf{X}_{ijk}\beta_3 + \delta_j + \gamma_k + \varepsilon_{ijk},$$

where D_{ijk} is equal to one if person i from country of origin j residing in area k receives disability payments and zero otherwise. Models are run separately for DI and SSI. We define area based on Public Use Microdata Areas (PUMAs).[6] The proportion of people receiving disability payments in a person's ethnic group is denoted $\overline{D}_j$.[7] This will refer to average DI take-up in DI

5. DDS award rates for DI applicants in the year 2000 ranged from 65% in New Hampshire to 31% in Texas. For SSI, they ranged from 59% in New Hampshire to 27% in West Virginia (Benitez-Silva, Buchinsky, and Rust 2004). It seems unlikely that these differences are attributable completely to differences in disability rates.

6. PUMAs are the smallest level of geography available in the 5% 2000 Census Public Use Micro Sample. They typically have about 100,000 residents. We also conducted the analysis measuring CA at the Metropolitan Statistical Area, and as can be seen in Tables S1 and S2 (available as Supporting Information in the online version of this article), results were similar. Dropping PUMAs in the five largest MSAs also yielded similar results (see columns 3 and 4 of Tables S1 and S2).

7. Another approach often used in the literature is to construct this average separately by PUMA. That might be a better measure of disability program take-up among the co-ethnics with which immigrants associate, but using such a variable may result in severe endogeneity bias. While people cannot choose average disability program take-up within their ethnic group across the entire country, they can choose this average in their PUMA through their residential choices. We also calculated take-up rates for co-ethnics living outside of the respondent's PUMA. In models with this measure, we used the $\overline{D}_{j,-k}$ as an additional control since it was no longer subsumed by the country of origin fixed effects. For clear comparison, we reran our baseline model using $\overline{D}_j$ as a control and dropping the country of origin fixed effects. For both DI and SSI, the estimated coefficients on the interaction variables were indistinguishable across the two model specifications.

models and average SSI take-up in SSI models. CA_{jk} refers to contact availability or the density of country of origin group j in area k. Contact availability is defined as $\log(C_{jk}/P_k)$, where C_{jk} is the number of people in area k who are from country of origin j and P_k is the population of area k.[8] Country of origin and area fixed effects are denoted δ_j and γ_k, respectively, while $\mathbf{X}_{ijk}$ is a vector of demographic characteristics including human capital, demographic, and assimilation variables. The country of origin fixed effects, area of residence fixed effects, and the contact availability control variable account for many omitted factors. The country of origin fixed effects also absorb the direct impact of $\overline{D}_j$. Our measure of networks will have the expected positive coefficient only if being surrounded by co-ethnics increases program participation more for people in ethnic groups with high disability program take-up.

A potential threat to our identification strategy is that immigrants who reside amid a large number of others with their ethnic background may be very similar to them in ways that can result in similar tendencies to participate in disability programs. For example, Cape Verdean immigrants residing in Cape Verdean enclaves may have characteristics which make them significantly more likely to find DI attractive than the Tanzanians who live in Tanzanian neighborhoods or other Cape Verdeans who do not live in Cape Verdean neighborhoods. To use the terminology of Manski (1993), a positive estimated coefficient in our interaction may simply reflect correlated effects which are a result of unobserved characteristics that affect individuals in a group simultaneously.[9] We take several steps to address potential threats to identification. First, we examine the effect of adding several country-of-origin–PUMA level variables to our baseline models. We also explore whether correlations in work histories can explain our DI results and correlations in poverty rates can explain our SSI results by examining retirees.

IV. DATA

Our source of data is the 5% sample of the 2000 U.S. Census as reported by the Integrated Public Use Microdata Series (IPUMS, Ruggles et al. 2010). Our sample consists of immigrants, age 25–61, who do not reside in group quarters and are not currently in school. Given the restrictions on SSI eligibility imposed by PRWORA in 1996, we limit our analysis to those immigrants who were living in the United States in 1995, 5 years prior to the survey. This restriction also increases the proportion of the sample eligible for DI payments given the program's work history conditions. Only naturalized citizens and noncitizens are considered immigrants, meaning that Puerto Ricans and people from other U.S. territories as well as individuals born abroad of American parents are dropped from the sample. In order to clearly differentiate between ethnicities, we drop American Indians, Alaskan natives, and Hawaiians from our sample and also individuals whose countries of origin are not clearly specified in the data. Finally, we merged Azores with Portugal, Korea with South Korea, and all countries comprising the United Kingdom together with each other in order to match the Census data with data from the WVS for our mechanisms analysis.

The U.S. Census does not directly ask whether people are receiving disability income. However, the Census does ask for the amount of income people are receiving from Social Security and SSI, separately. Technically, Social Security income can be in the form of disability insurance as well as public pensions, survivor benefits, and Railroad Retirement insurance payments, but it is unlikely that people in our sample are receiving pensions given that they are all below even the early retirement age. We also drop widows and widowers from the sample to make it less likely that they are receiving survivor benefits.[10] Similarly, SSI payments can be made to the disabled as well as the elderly, but given the age

8. We use a log specification both because of the tremendous variation in contact availability in the data and because this is what is typically used in the literature. As can been seen in Table S3 (columns 3 and 6) our results are robust to dropping the log. Results are also robust to following Bertrand et al.'s specification of contact availability (columns 2 and 5). They weight the proportion of the PUMA's population that is of the person's ethnic group by the ratio of the total number of people in the country who belong to the ethnic group to the total number of people in the country.

9. In contrast, endogenous effects occur when individual behaviors vary causally with the behaviors of group members and exogenous effects occur when individual behaviors vary causally with exogenous attributes of group members. We will not be able to distinguish endogenous from exogenous effects but we will examine the likely mechanisms through which network effects operate in Sections VI and VII.

10. Of the 11,280,792 DI recipients in 2010, only 160,300 were receiving spouse benefits and 97,518 were receiving benefits as disabled adult children of disabled workers (Annual Statistical Report on the Social Security Disability Insurance Program 2010). Using our sample of immigrants, results were robust to dropping households with more than one disability payment recipient.

TABLE 2
Descriptive Statistics

	Whole Sample		DI Sample	SSI Sample
	(1) Mean	(2) Standard Deviation	(3) Mean	(4) Mean
DI	0.017	0.127	—	0.154
SSI	0.013	0.113	0.121	—
Age	41.03	9.658	47.05	46.363
Male	0.508	0.499	0.520	0.459
High school dropout	0.321	0.466	0.446	0.541
High school degree	0.303	0.459	0.313	0.295
Some college	0.146	0.353	0.121	0.092
English fluency	0.487	0.499	0.410	0.334
Spouse present	0.687	0.463	0.602	0.475
Child	0.642	0.479	0.578	0.533
Number of children	2.212	1.189	2.147	2.352
Hispanic	0.226	0.418	0.232	0.209
Black	0.075	0.263	0.091	0.091
Asian	0.248	0.431	0.169	0.250
Other race	0.004	0.064	0.003	0.004
Years in the United States	18.63	10.404	22.773	20.310
Disability	0.175	0.379	0.309	0.425
Contact availability (CA) in levels	0.068	0.101	0.077	0.079
CA	−4.172	2.062	−4.035	−3.897
Observations	704,871		11,955	9,314

Notes: All observations in our sample (described in the notes to Table 1) are used to construct the statistics in Columns 1 and 2. The sample is restricted to DI recipients in column 3 and to SSI recipients in column 4. DI is a dummy variable that equals one if a person receives disability insurance income. SSI is a dummy variable that equals one if a person receives Supplemental Security Income. Child is a dummy variable that equals one if the person has at least one child living in the household whereas "English fluency" equals one for people who speak "only English at home" or speak English "very well" and zero for people who speak "well," "not well," or "not at all." Disability is a dummy variable signifying a lasting physical or mental health condition that causes difficulty working, limits the amount or type of work the person can do, or prevents him or her from working altogether. CA, contact availability, is the log of the proportion of people residing in the PUMA that are from the person's country of origin. CA was calculated using all observations in the 2000 5% Census extract (14.1 million observations). Estimates are weighted using the appropriate person-level weights provided by the 2000 U.S. Census.

restrictions we impose on the data, recipients of SSI in our sample would be receiving it as a result of a disability. Our final sample consists of 704,871 observations.

Table 2 shows descriptive statistics of the variables used in the analysis. The proportions of our sample that receive DI and SSI are similar. This pattern differs from the general population where, among those receiving payments on the basis of a disability, over twice as many people receive DI alone than SSI alone (Chart 12, Annual Statistical Report on the DI Program 2010). We remind readers that the foreign born are significantly less likely to satisfy the DI work history requirements because they may not have resided in the United States for a sufficient number of years and also they are more likely to work "under the table" or not work at all in the years they have resided in the United States. Given their typical lower earnings than natives (Larsen 2004), immigrants are also more likely to qualify for SSI. For further details on how immigrants compare to natives in terms of SSI receipt, see Kaushal (2010) which examines elderly immigrants' labor supply responses to changes in SSI requirements in 1996.

Table 2 also shows that on average, disability payment recipients are older, have lower levels of education, and are more likely to live in PUMAs with a large representation of co-ethnics. Immigrants in our sample have lived in the United States 18.6 years on average, making them very likely to be eligible for DI. In fact, 80% of our sample have lived in the United States for more than 10 years. Racial distributions do not differ substantially by whether people participate in disability programs. Comparing DI recipients to SSI recipients, we can see that DI recipients have higher levels of education and English fluency than SSI recipients. DI recipients typically have resided in the United States for a longer period of time. Asians are significantly more likely to receive SSI than DI. Beyond these differences, DI and SSI recipients have very similar observable characteristics. Some immigrants in our sample receive disability payments from both DI and

TABLE 3
Effects of Networks on DI/SSI Receipt

	DI		SSI	
	(1)	(2)	(3)	(4)
CA * proportion of co-ethnics receiving DI	0.098**	0.094**		
	(0.024)	(0.024)		
CA * proportion of co-ethnics receiving SSI			0.269**	0.270**
			(0.030)	(0.030)
CA	−0.001**	−0.001**	−0.003**	−0.004**
	(0.0004)	(0.0004)	(0.0004)	(0.0004)
Male	0.001**	0.0004	−0.002**	−0.003**
	(0.0003)	(0.0003)	(0.0003)	(0.0003)
Hispanic	0.001***	0.001***	−0.001***	−0.001*
	(0.0005)	(0.005)	(0.0005)	(0.0004)
Black	0.007**	0.007**	0.005**	0.003*
	(0.002)	(0.002)	(0.002)	(0.002)
Asian	−0.0004	0.001	−0.001	−0.001
	(0.001)	(0.001)	(0.001)	(0.001)
Other race	−0.003	−0.003	−0.001	−0.003
	(0.002)	(0.002)	(0.003)	(0.003)
High school dropout		0.012**		0.018**
		(0.001)		(0.001)
High school degree		0.006**		0.008**
		(0.0004)		(0.0004)
Some college		0.003**		0.003**
		(0.0005)		(0.0003)
Spouse present		−0.008**		−0.013**
		(0.0005)		(0.0005)
Child		0.001***		−0.001*
		(0.0006)		(0.0005)
Number of children		−0.001*		0.0002
		(0.0002)		(0.0002)
English fluency		−0.003**		−0.002**
		(0.0004)		(0.0004)
Disability		0.013**		0.020**
		(0.001)		(0.0006)
Observations	704,871	704,871	704,871	704,871
Adjusted R^2	0.013	0.017	0.018	0.028
Country of origin fixed effects	Yes	Yes	Yes	Yes
PUMA fixed effects	Yes	Yes	Yes	Yes
Age fixed effects	Yes	Yes	Yes	Yes
Years in the U.S. fixed effects	No	Yes	No	Yes

Notes: See Table 1 notes for information on the sample and Table 2 for notes on the variables. In columns 1 and 2 the dependent variable is Disability Insurance (DI) and in columns 3 and 4 the dependent variable is Supplemental Security Income (SSI). Coefficients are estimated using linear probability models. The omitted education dummy is "College and more." The omitted race dummy is "white." Heteroskedasticity corrected standard errors clustered by geography and country of origin cells are in parentheses. Estimates are weighted using the appropriate person-level weights provided by the 2000 U.S. Census.

*Significant at 10%; **significant at 5%; ***significant at 1%.

SSI—12.1% of DI recipients receive SSI and 15.4% of SSI recipients receive DI.

The CA variable suggests that on average immigrants in our sample live in PUMAs where 7% of the population shares their country of origin. About 25% of the immigrants in our sample live in PUMAs where less than 0.3% of the population is from their country of origin while a little over 5% live in PUMAs where more than 30% of the population shares their ethnic origin.

V. EMPIRICAL RESULTS

A. Baseline Results

Table 3 present estimates of the coefficients in Equation (1) for linear probability models explaining DI and SSI participation.[11]

11. We also estimated this equation using logit and probit models without PUMA fixed effects. As can be seen in Table S4, marginal effects of our network measure were positive and significant at the 1% level.

Our parameters of interest are identified from variation across 130 countries of origin and 2,071 PUMAs. Standard errors are clustered on country-of-origin-PUMA cells, but results are robust to clustering either on country of birth or PUMA individually. As can be seen in the first and third columns, our estimates in both the DI and SSI models suggest a positive and statistically significant coefficient on the interaction between contact availability and the proportion of co-ethnics receiving disability program payments, even in very simple models containing only the controls that can be considered reasonably exogenous. The estimated CA coefficients are negative suggesting that living in ethnic enclaves actually decreases the probability of disability program take-up for immigrants belonging to ethnic groups with lower rates of take-up. Males are more likely to have substantial work histories and hence they are more likely to receive DI than females but less likely to receive SSI. Blacks are more likely to receive disability payments of both types than the other racial groups. Compared to whites, Hispanics are less likely to receive SSI but are marginally more likely to receive DI.

In the second and fourth columns, additional demographic and human capital controls—including schooling, years in the U.S. fixed effects, and whether the person has a work-preventing disability—are added. Naturally, people with disabilities are more likely to receive disability benefits.[12] Married people are less likely to receive both types of disability payments. Immigrants with more education and better English-speaking abilities are less likely to be receiving DI and SSI. Most interestingly, when these variables are added to both the DI and SSI models, the estimated network coefficients do not change substantially.[13] This suggests that the country of origin and PUMA fixed effects are likely to be controlling for the most influential unobservable characteristics.

Our final DI model suggests that for an immigrant in an ethnic group with average DI take-up (0.017), a 10% increase in the proportion of co-ethnics increases the likelihood of going on DI by a statistically insignificant 0.006 percentage points. On the other hand, for immigrants in an ethnic group with take-up rates of 0.046, the highest take-up rate in our sample, the same 10% increase in the proportion of co-ethnics increases the likelihood of going on DI by a statistically significant 0.033 percentage points. These numbers imply that living amid co-ethnics increases DI take-up over five times more for immigrants in groups with the highest DI take-up than immigrants in groups with average DI take-up. There is even more variation in sensitivity to exposure to co-ethnics in SSI take-up. Our final SSI model implies that among immigrants in groups with average SSI take-up (0.013), a 10% increase in the proportion of co-ethnics actually results in a statistically insignificant 0.005 percentage point decrease in the probability of SSI take-up. However, the same increase in exposure to co-ethnics leads to a statistically significant 0.211 percentage point increase in the probability of take-up for those in groups with the highest SSI take-up, 0.093.

It may not be surprising that network effects are stronger for SSI take-up than those of DI take-up for two reasons. First, person-to-person information sharing should be relatively more important for people eligible for SSI payments given their low life-time earnings and lower levels of human capital. Second, while DI is an insurance program requiring recipients to have paid into Social Security, SSI is a means-tested program. This may suggest that any taboos against exaggerated disability claims are more important for SSI than for DI.

B. Robustness Checks

The main potential threat to our identification strategy is the possibility that immigrants who choose to reside amid a large number of co-nationals are more similar to the people in their ethnic groups in unobservable ways which then results in similarities in disability program participation. A particular concern is that immigrants residing amid a large number of other immigrants from their country of origin are likely to be employed in the same types of jobs. The Census contains information on people's occupation and industry, but only for people who have worked within the previous 5 years. The disabled are typically no longer employed, and when they are, it is unlikely that they still have the job from which they were laid off or that caused their disability. Thus, it is not straightforward to control people's

12. Admittedly, this is a rather crude measure of disability. However, in a companion paper using NHIS data (Furtado and Theodoropoulos 2013), we find network effects to be robust to models controlling for health behaviors, such as smoking, and more subjective measures of general health.

13. In fact, even when we allow the effects of these controls to differ by country of origin, the estimated network coefficients fall in magnitude but remain qualitatively the same (results available upon request).

listed occupations and industries.[14] However, we construct several aggregate variables which can be used to alleviate the most obvious occupation-related concerns with our identification strategy.

We start by considering the role of labor markets. Using plausibly exogenous variation resulting from coal booms and busts, Black, Daniel, and Sanders (2002) find that economic conditions have strong impacts on both DI and SSI participations. Plant downsizing in Norway has also been found to substantially increase disability program participation of workers in affected plants (Rege, Telle, and Votruba 2009). To explore whether labor market opportunities are driving our results, we construct country-of-origin-PUMA unemployment rates and mean log wages and examine whether adding these variables has any impact on the estimated network coefficients.

Another issue we consider is whether immigrants from high disability program participation groups living in ethnic enclaves are especially likely to receive disability payments simply because they are more likely to have become injured on the job. Starting with data from the Bureau of Labor Statistics' (BLS) Injuries, Illnesses, and Fatalities program on work-related fatalities and nonfatal injuries and illnesses in 2003–2005, we follow Orrenius and Zavodny (2009) in constructing on-the-job injury and fatality rates. Specifically, we divide the number of injuries or fatalities in the occupation by the number of private sector workers in the occupation.[15] Data on the number of workers in each occupation are obtained from the occupational employment statistics. After assigning to each employed person in the full sample injury rates for his or her occupation, we then construct average injury rates for each PUMA-country-of-origin cell. We do the same for fatality rates and explore whether adding these variables to the model changes our estimated network coefficients. Descriptive statistics on all of these aggregate variables are shown in Table S5.

Tables 4 and 5 present results from DI and SSI models that include controls for labor market conditions as well as occupational hazards. The first columns show that results from our baseline model run on individuals with non-missing data on the aggregate variables are almost identical to results from the full sample.[16] As seen in the second columns of both tables, aggregate unemployment rates are positively associated with disability program take-up while wages are negatively correlated. Strangely, immigrants residing in areas where people from their country of origin tend to work in jobs with high injury are less likely to receive SSI. In any case, adding these aggregate labor market characteristics variables to our models has no impact on the estimated network coefficients.

Next, we consider impacts of aggregate-level versions of several individual-level variables we control in our baseline regressions (again, see Table S5 for descriptive statistics). In the third columns of Tables 4 and 5, we add average values of years of schooling, years in the United States, and age in a person's PUMA-country-of-origin cell to the models. These aggregate variables do have an impact on disability program take-up, even when controlling for the individual level versions of these variables. Nevertheless, their inclusion does not change our estimated network effects. In column 4, we present results of models controlling for all aggregate variables, and results remain the same. We conclude from these analyses that estimated network effects in these models are very robust.

To alleviate any concerns that our results may be driven by functional form or outliers, we then performed simple ratios of ratios and differences in differences exercises. Results, shown in Table S6, show that the basic relationships we find in the data are not sensitive to our choice of controls and that the same story can be told whether effects are additive or multiplicative.

VI. SOCIAL NORMS AND DISABILITY PROGRAM TAKE-UP

A. Cheating the Government Taboos and Importance of Work Norms

Having provided evidence that social interactions play an important role in immigrants'

14. We did run regressions with occupation fixed effects for both DI and SSI (results upon request). Estimated network coefficients were practically the same as those in our baseline model for DI but significantly smaller than the baseline for SSI. We remain cautious about interpreting these results since only 35% of SSI recipients list an occupation in the Census; 66% of DI recipients list an occupation.

15. A work-related injury is defined as an injury involving at least one full day away from work. Occupations with the highest injury rates are farmers and ranchers, fishers and hunters, loggers, and mining machine operators.

16. There were some PUMA-country-of-origin cells containing only individuals who do not list an occupation or who list an occupation for which we do not have data on occupational hazards because they are self-employed, for example.

TABLE 4
Effects of Aggregate Variables on DI

Dependent Variable: DI	(1)	(2)	(3)	(4)
CA * proportion of co-ethnics receiving DI	0.114*	0.114*	0.118*	0.118*
	(0.024)	(0.024)	(0.024)	(0.024)
CA	−0.001*	−0.001*	−0.002*	−0.002*
	(0.0004)	(0.0004)	(0.0004)	(0.0004)
Unemployment rate within PUMA-country of origin		0.004		0.004
		(0.003)		(0.003)
Log of average yearly wage income within PUMA-country of origin		−0.001**		−0.0004
		(0.0005)		(0.0005)
On-job injury rates within PUMA-country of origin		−0.0004		−0.0005
		(0.0004)		(0.0005)
On-job fatality rates within PUMA-country of origin		0.0003		0.0003
		(0.0004)		(0.0004)
Average years of schooling within PUMA-country of origin			−0.012	−0.011
			(0.013)	(0.014)
Average years in the United States within PUMA-country of origin			−0.021*	−0.021*
			(0.005)	(0.005)
Average age within PUMA-country of origin			−0.026*	−0.026*
			(0.006)	(0.006)
Observations	682,378	682,378	682,378	682,378
Adjusted R^2	0.016	0.016	0.016	0.016
Country of origin fixed effects	Yes	Yes	Yes	Yes
PUMA fixed effects	Yes	Yes	Yes	Yes
Age fixed effects	Yes	Yes	Yes	Yes
Years in the U.S. fixed effects	Yes	Yes	Yes	Yes
Controls	Yes	Yes	Yes	Yes

Notes: See Table 1 notes for information on the sample and Table 2 for notes on the control variables. Coefficients are estimated using linear probability models. For reasons of brevity, we only present coefficients of the aggregate variables. Other controls are those shown in Table 2. Heteroskedasticity corrected standard errors clustered by geography and country of origin cells are in parentheses. The aggregate injury, fatality, schooling, years in the United States, and age variables are divided by 100. Estimates are weighted using the appropriate person-level weights provided by the 2000 U.S. Census.

*Significant at 10%; **significant at 5%.

disability program take-up, in this section we explore why. Knowing the mechanisms through which networks operate is particularly important from a policy perspective because while some types of social interactions generate multiplier effects (endogenous effects), others do not (exogenous effects). Although we are not able to perfectly distinguish between the mechanisms driving our network results, in this section we present evidence suggesting that social norms, an often-cited source of endogenous effects, play some role in explaining network effects in disability program participation.

While exaggerating a disability in order to receive benefits may be stigmatized in certain ethnic communities, it may be less taboo or even admired in others. We may expect then that exposure to disability income recipients increases take-up more among people belonging to groups with more lax taboos against receiving benefits despite having only a minor disability. In addition, because receiving benefits usually implies leaving the labor force, norms may operate via people's beliefs about the importance of work. Even the severely disabled may continue to work despite significant hardship in order to preserve a sense of dignity in communities with strong work norms. Thus, exposure to disability program participants may increase the probability of take-up less for those in groups with strong work norms.

To gain insight into the role of social norms in driving ethnic network effects, we turn to data from the European and World Values Surveys (WVS) four-wave integrated data file (European and World Values Survey Association 1981–2004), a compilation of national surveys on a variety of topics including attitudes toward cheating the government and the importance of work. Starting with individual-level data from the WVS, we construct aggregate measures of home country norms which we then merge with our Census sample by country of origin. We generally use data from the 2000 wave of the WVS, but if a question was not asked in a country in the 2000 wave but asked in the 1995 wave, we used

TABLE 5
Effects of Aggregate Variables on SSI

Dependent Variable: SSI	**(1)**	**(2)**	**(3)**	**(4)**
CA * proportion of co-ethnics receiving SSI	0.302**	0.301**	0.303**	0.301**
	(0.031)	(0.031)	(0.031)	(0.031)
CA	−0.004**	−0.004**	−0.004**	−0.004**
	(0.0004)	(0.0004)	(0.0004)	(0.0004)
Unemployment rate in PUMA-country of origin		0.006*		0.006*
		(0.003)		(0.003)
Log of yearly wage income in PUMA-country of origin		−0.0006		0.0003
		(0.0005)		(0.0005)
On-job injury rates in PUMA-country of origin		−0.001**		−0.002**
		(0.0004)		(0.0004)
On-job fatality rates in PUMA-country of origin		0.001		0.001
		(0.001)		(0.001)
Average years of schooling in PUMA-country of origin			−0.043**	−0.057**
			(0.012)	(0.014)
Average years in the United States in PUMA-country of origin			−0.018**	−0.019**
			(0.005)	(0.005)
Average age in PUMA-country of origin			−0.021**	−0.020**
			(0.006)	(0.006)
Observations	682,378	682,378	682,378	682,378
Adjusted R-squared	0.028	0.028	0.028	0.028
Country of origin fixed effects	Yes	Yes	Yes	Yes
PUMA fixed effects	Yes	Yes	Yes	Yes
Age fixed effects	Yes	Yes	Yes	Yes
Years in the U.S. fixed effects	Yes	Yes	Yes	Yes
Controls	Yes	Yes	Yes	Yes

Notes: See Table 1 notes for information on the sample and Table 2 for notes on the control variables. Coefficients are estimated using linear probability models. For reasons of brevity, we only present coefficients of the aggregate variables. Other controls are those shown in Table 2. Heteroskedasticity corrected standard errors clustered by geography and country of origin cells are in parentheses. The aggregate injury, fatality, schooling, years in the United States, and age variables are divided by 100. Estimates are weighted using the appropriate person-level weights provided by the 2000 U.S. Census.

*Significant at 10%; **significant at 5%.

the 1995 responses. In the end, we were able to match all our WVS variables of interest to 37 of the 130 countries in our Census sample.[17]

We used several questions from the WVS to gain insight into home country norms. For example, one of our measures of the government cheating taboo uses a question on whether people think cheating on taxes can always be justified, never be justified, or something in between these two extremes (the scale runs from 1 to 10). From the individual responses to these questions, we construct a variable measuring the proportion of a person's home country that believes these actions are "Never Justified." Similar measures were constructed from questions asking whether claiming government benefits for which one is not eligible is ever justified and whether taking public transport without paying the fare is ever justified.

To measure work norms, we constructed variables measuring the percentage of people in a person's home country strongly agreeing that work should come first (even if it means less spare time), that work is a duty towards society, that in order to develop talents you need to have a job, and that people who do not work turn lazy. We also constructed a variable measuring the percentage of people saying that work is "very important" in life (as opposed to rather important, not very important, or not at all important) and another one measuring the percentage of people believing that compared to leisure, work is what makes life worth living. Descriptive statistics on all of these variables are shown in Table S7.

17. Our sample size falls by 43% when limited to observations with non-missing data on all of our WVS variables. When running our baseline model on this smaller sample, the baseline network coefficient in the DI model is not precisely estimated ($p = .13$, see column 1 of Table 6). However, the magnitude is roughly the same as that shown in column 2 of Table 3, and the two estimates are statistically indistinguishable. For SSI, network effects estimated using both samples are statistically significant and very similar to each other (see column 1 of Table 7 and column 4 of Table 3). To check for robustness, we replaced our missing WVS data with zeros and then added dummy variables for missing data interacted with our variables of interest. A similar story unfolds. These results are available upon request.

Because there are several WVS questions essentially measuring the same concepts, we aggregated information using principal components analysis. As can be seen in Table S8, the eigenvalues of 2.4 and 3.2 for the first component of the government cheating taboos and work norms, respectively, far exceed the rule of thumb number of one. Moreover, the first principal component explains more than half of the common variance of the three measures of government cheating taboos and over 80% of the common variance of the five measures of work norms. All factors load positively on both first principal components. For all these reasons, we aggregate the government cheating taboo and importance of work variables by constructing the first principal component of each[18] and estimate equations with the following basic form:

$$(2)\quad D_{ijk} = \gamma_1 \overline{\mathrm{Norm}}_j \times \overline{D}_j \times \mathrm{CA}_{jk} + \gamma_2 \mathrm{CA}_{jk} + \gamma_3 \overline{D}_j \times \mathrm{CA}_{jk} + \gamma_4 \overline{\mathrm{Norm}}_j \times \mathrm{CA}_{jk} + \mathbf{X}_{ijk}\gamma_5 + \delta_j + \gamma_k + u_{ijk},$$

where $\overline{\mathrm{Norm}}_j$ takes on higher values when countries have norms which make it more "costly" to take-up benefits either as a result of strong government cheating taboos or importance of work norms. All other variables are defined as before. If network effects operate via social norms, we expect γ_1 to be negative since exposure to co-ethnics receiving disability payments should result in relatively less take-up among people from countries with stronger government cheating taboos or importance of work norms.[19]

Columns 1 through 4 of Tables 6 and 7 present results for DI and SSI, respectively.[20] Only estimated coefficients on the triple and double interactions are shown in the tables, but the full set of the original control variables and fixed effects are included in the models. Results suggest that both cheating the government taboos and importance of work norms decrease estimated network effects for DI when they are included in the model individually. However, while they remain of roughly the same magnitudes, only the estimated importance of work triple interaction coefficient is statistically significant when both triple interactions are included in the model at the same time. Government cheating taboos and work norms have similar (negative) impacts on SSI-related network effects when included in models individually, but neither triple interaction estimated coefficient is statistically significant when both triple interactions are included in the model at the same time. These estimates suggest that norms do play a role in explaining network effects, but we cannot say anything conclusively about the influence of government cheating taboos relative to importance of work norms. Both seem to have important impacts on DI and SSI take-up.

A simple example can help clarify the interpretation of the estimates of the coefficients on the triple interactions. Consider the estimate of −0.08 on the importance of work norms triple interaction in the DI model shown in column 3 of Table 6. Take two immigrants from ethnic groups with the same average DI take-up (0.015) and living in a PUMA with the same representation of their ethnic groups but from countries with different views on the importance of work. A 10% increase in the share of co-ethnics living in the same PUMA would result in a 0.039 percentage point increase in the likelihood of receiving DI for the immigrant from a country at 25th percentile of our measure of importance of work norms. The same increase in co-ethnic share would result in a virtually zero percentage point increase (0.0000125) in the probability of DI participation for an immigrant from a country at the 75th percentile.

We also ran similar models using the individual variables constructed from the WVS. Findings, shown in Tables S9 and S10, suggest that our results are robust across different measures of home country norms. With few exceptions, the estimated triple interaction coefficients are negative and statistically significant in both DI and SSI models.

B. *Other Potential Mechanisms*

Although we view these results as suggestive of the role of social norms in explaining network effects, we cannot conclude that norms are the only drivers of network effects. Information sharing regarding the existence of the

18. We also aggregated variables using simple averages, and results were similar.

19. We were also interested in whether exposure to co-ethnics directly leads to lower disability program take-up among immigrants from countries with stronger government cheating taboos and work norms. Interestingly, the estimated coefficients on the interactions between contact availability and our norms variables were not robust and usually statistically insignificant. We conclude, therefore, that norms can exacerbate or attenuate immigrants' reactions to exposure to disability program take-up in their ethnic communities, but norms in themselves do not have strong or consistent impacts on take-up of these programs.

20. We describe the results in columns 5–8 of these tables in Section VII.

TABLE 6
Social Norms and Social Security-Baseline and Retirement-Age Samples

	Baseline Sample Dependent Variable: DI				Retirement-Age Sample Dependent Variable: Social Security			
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
CA * proportion of co-ethnics receiving DI (col 1–4) or Social Security retirement income (col 5–8)	0.120 (0.080)	0.137 (0.087)	0.053 (0.070)	0.041 (0.078)	0.029* (0.013)	0.039** (0.014)	0.046** (0.014)	0.045** (0.014)
CA	−0.001 (0.001)	−0.002 (0.002)	−0.001 (0.001)	−0.0004 (0.001)	−0.027** (0.008)	−0.034** (0.009)	−0.039** (0.009)	−0.038** (0.009)
CA * proportion of co-ethnics receiving DI (col 1–4) or Social Security retirement income (col 5–8) * PCA cheating government taboos		−0.093*** (0.048)		−0.076 (0.059)		−0.013 (0.008)		0.004 (0.009)
CA * PCA cheating government taboos		0.001*** (0.0006)		0.001 (0.001)		0.006 (0.004)		−0.001 (0.004)
CA * proportion of co-ethnics receiving DI (col 1–4) or Social Security retirement income (col 5–8) * PCA work norms			−0.080** (0.028)	−0.074* (0.032)			−0.008 (0.007)	−0.008 (0.007)
CA * PCA work norms			0.001* (0.0003)	0.001*** (0.0004)			0.009*** (0.005)	0.010*** (0.005)
Observations	403,288	403,288	403,288	403,288	50,313	50,313	50,313	50,313
Adjusted R^2	0.015	0.015	0.015	0.015	0.302	0.302	0.302	0.302
Years in the U.S. fixed effects	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Country of origin fixed effects	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
PUMA fixed effects	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Age fixed effects	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Controls	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes

Notes: Columns 1 through 4 use the baseline sample (25–61 years old) and columns 5 through 8 use the retirement-age sample (65 and above). Due to the missing observations in the WVS raw data, the coefficients in columns 1 through 4 (5 through 8) are estimated using 403,288 (50,313) observations as opposed to 704,871 (100,090) observations in the baseline (retirement-age) sample(s). Column 1 estimates our baseline specification (i.e., column 2 of Table 3) but using 403,288 observations. Column 5 estimates our baseline specification (i.e., column 2 of Table 3) on the retirement-age sample but using 50,313 observations. In columns 1 through 4, networks are measured by the interaction between CA and the proportion of co-ethnics receiving DI while in columns 5 through 8, they are measured by the interaction between CA and the proportion of co-ethnics receiving Social Security retirement income. The "PCA cheating government taboos" is the first principal component constructed using the following three variables: "percent of co-ethnics saying cheating on taxes is never justifiable," "percent of co-ethnics saying claiming government benefits to which not entitled is never justifiable," and "percent of co-ethnics saying avoiding fare on public transport is never justifiable." The "PCA work norms" is the first principal component constructed using the following six variables: "percent of co-ethnics saying work is very important in life," "percent of co-ethnics saying work is what makes life worth living," "percent of co-ethnics strongly agreeing that people who do not work turn lazy," "percent of co-ethnics strongly agreeing that to develop talents you need to have a job," "percent of co-ethnics strongly agreeing that work is a duty toward society," and "percent of co-ethnics strongly agreeing that work should come first even if it means less spare time." In columns 2 and 6 we include interactions between "PCA cheating government taboos" and our variables of interest. In columns 3 and 7 we include interactions between "PCA work norms" and our variables of interest, and in columns 4 and 8, we include interactions between both PCA variables and our variables of interest. For reasons of brevity, we only report coefficients on the variables of interest. Other controls include those shown in Table 2. Coefficients are estimated using linear probability models. Heteroskedasticity corrected standard errors clustered by geography and country of origin cells are in parentheses. Estimates are weighted using the appropriate person-level weights provided by the 2000 U.S. Census.

*Significant at 10%; **significant at 5%; ***significant at 1%.

programs, how to complete the necessary paperwork, the doctors who provide the most convincing cases for disability, and the lawyers who are most successful in appealing negative decisions may also generate network effects. The literature suggests that the importance of information sharing in explaining program take-up depends on the particular program and context. Some papers find strong roles of information sharing (Aizer 2007; Figlio, Hamersma, and Roth 2011) and others find no evidence of information sharing (Aizer and Currie 2004; Åslund and Fredriksson 2009). Unfortunately, we do not have a clean way to test for information sharing with

TABLE 7
Social Norms and SSI-Baseline and Retirement-Age Samples

	Baseline Sample Dependent Variable: SSI				Retirement-Age Sample Dependent Variable: SSI			
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
CA * proportion of co-ethnics receiving SSI	0.253** (0.032)	0.266** (0.036)	0.213** (0.034)	0.176* (0.080)	0.140** (0.017)	0.135** (0.017)	0.147** (0.018)	0.146**
CA	−0.003** (0.0004)	−0.003** (0.0004)	−0.002** (0.0004)	−0.002* (0.001)	−0.012** (0.003)	−0.011** (0.003)	−0.012** (0.003)	−0.012** (0.003)
CA * proportion of co-ethnics receiving SSI * PCA cheating government taboos		−0.032* (0.014)		0.007 (0.035)		−0.013*** (0.008)		0.0004 (0.009)
CA * PCA cheating government taboos		0.001** (0.0001)		0.0002 (0.0004)		0.003*** (0.002)		−0.001 (0.002)
CA * proportion of co-ethnics receiving SSI * PCA work norms			−0.046* (0.018)	−0.074 (0.046)			−0.007 (0.011)	−0.005 (0.012)
CA * PCA work norms			0.0003*** (0.0002)	0.0001 (0.0003)			−0.003 (0.001)	−0.003*** (0.001)
Observations	403,288	403,288	403,288	403,288	50,313	50,313	50,313	50,313
Adjusted R^2	0.302	0.302	0.302	0.302	0.222	0.222	0.222	0.222
Years in the U.S. fixed effects	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Country of origin fixed effects	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
PUMA fixed effects	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Age fixed effects	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Controls	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes

Notes: Columns 1 through 4 refer to the baseline sample (25 to 61 years old) and columns 5 through 8 refer to the retirement-age sample (65 and above). Due to the missing observations in the WVS raw data, the coefficients in columns 1 through 4 (5 through 8) are estimated using 403,288 (50,313) observations as opposed to 704,871 (100,090) observations in the baseline (retirement-age) sample(s). Column 1 estimates our baseline specification (i.e., column 2 of Table 3) but using 403,288 observations. Column 5 estimates our baseline specification (i.e., column 2 of Table 3) in the retirement-age sample but using 50,313 observations. In columns 1 through 4, networks are measured by the interaction between CA and the proportion of *working-age* co-ethnics receiving SSI while in columns 5 through 8, they are measured by the interaction between CA and the proportion of *retirement-age* co-ethnics receiving SSI. The "PCA cheating government taboos" is the first principal component constructed using the following three variables: "percent of co-ethnics saying cheating on taxes is never justifiable," "percent of co-ethnics saying claiming government benefits to which not entitled is never justifiable," and "percent of co-ethnics saying avoiding fare on public transport is never justifiable." The "PCA work norms" is the first principal component constructed using the following six variables: "percent of co-ethnics saying work is very important in life," "percent of co-ethnics saying work is what makes life worth living," "percent of co-ethnics strongly agreeing that people who do not work turn lazy," "percent of co-ethnics strongly agreeing that to develop talents you need to have a job," "percent of co-ethnics strongly agreeing that work is a duty towards society," and "percent of co-ethnics strongly agreeing that work should come first even if it means less spare time." In columns 2 and 6 we include interactions between "PCA cheating government taboos" and our variables of interest. In columns 3 and 7 we include interactions between "PCA work norms" and our variables of interest, and in columns 4 and 8, we include interactions between both PCA variables and our variables of interest. For reasons of brevity, we only report coefficients on the variables of interest. Other controls include those shown in Table 2. Coefficients are estimated using linear probability models. Heteroskedasticity corrected standard errors clustered by geography and country of origin cells are in parentheses. Estimates are weighted using the appropriate person-level weights provided by the 2000 U.S. Census.

*Significant at 10%; **significant at 5%; ***significant at 1%.

our data. As a suggestive test only, we explore whether network effects are stronger for people with less formal education under the presumption that those with college degrees are able to gather information about the disability programs without requiring information from ethnic networks. Table 8 presents results from models including interactions of our network measure and educational attainment. Columns 1 and 3 show some evidence in favor of information sharing in explaining SSI take-up but no evidence of information sharing in DI take-up. High school dropouts and graduates are significantly more sensitive to exposure to SSI take-up within their ethnic groups than college graduates, but there is no difference across education levels in terms of sensitivity to DI exposure. For DI, none of the estimated triple interactions coefficients between

TABLE 8
Information Sharing and Leisure Complementarities, DI and SSI

	DI		SSI	
	(1)	(2)	(3)	(4)
CA * proportion of co-ethnics receiving DI	0.043	0.112**		
	(0.033)	(0.025)		
CA * proportion of co-ethnics receiving SSI			0.099**	0.307**
			(0.025)	(0.030)
CA	0.0001	0.0003	−0.001**	0.002**
	(0.0004)	(0.001)	(0.0003)	(0.001)
CA * proportion of co-ethnics receiving DI * HS dropout	0.045			
	(0.062)			
CA * proportion of co-ethnics receiving DI * HS degree	0.017			
	(0.043)			
CA * proportion of co-ethnics receiving DI * some college	0.026			
	(0.043)			
CA * proportion of co-ethnics receiving SSI * HS dropout			0.307**	
			(0.053)	
CA * proportion of co-ethnics receiving SSI * HS degree			0.068*	
			(0.032)	
CA * proportion of co-ethnics receiving SSI * some college			0.002	
			(0.029)	
CA * proportion of co-ethnics not working		−0.006**		−0.019**
		(0.002)		(0.002)
Observations	704,871	704,871	704,871	704,871
Adjusted R^2	0.017	0.017	0.032	0.029
Country of origin fixed effects	Yes	Yes	Yes	Yes
PUMA fixed effects	Yes	Yes	Yes	Yes
Age fixed effects	Yes	Yes	Yes	Yes
Years in the U.S. fixed effects	Yes	Yes	Yes	Yes
Controls	Yes	Yes	Yes	Yes

Notes: In columns 1 and 2 the dependent variable is Disability Insurance (DI) and in columns 3 and 4 the dependent variable is Supplemental Security Income (SSI). The triple interaction coefficients in column 1 are jointly not statistically different from zero (F statistic = 0.28, p value = .841), whereas they are jointly statistically different from zero in Column 3 (F statistic= 12.00, p value < .001). The variable "Proportion of co-ethnics not-working" is the proportion of co-ethnics who are either unemployed or out of the labor force. For reasons of brevity, we only report coefficients on the variables of interest. Other controls include all of the necessary double interactions (for columns 1 and 3 only) as well as the controls shown in Table 2. Estimates are weighted using the appropriate person-level weights provided by the 2000 U.S. Census.

*Significant at 10%; **significant at 5%.

education level, contact availability, and average DI take-up are statistically different from zero; we cannot reject the null hypothesis that the coefficients are jointly equal to zero. This may not be surprising given that the information provided through networks may be especially useful for the population qualifying for SSI benefits but not DI benefits. We note, however, that these patterns might also be explained by social norms if low education potential SSI recipients are especially sensitive to social norms within their ethnic groups.

We also explore whether complementarities in leisure are driving our network results. If the main reason people are more likely to take up disability programs when they are surrounded by others on these programs is that the availability of non-working friends makes leisure more enjoyable, then being surrounded by others who are out of the labor force for reasons unrelated to disability should have similar impacts on disability program take-up. We add to our baseline models an interaction between contact availability and the percentage of co-ethnics, living throughout the country, that are not employed. As can be seen in columns 2 and 4 of Table 8, the estimated coefficients on the not employed-contact availability interactions are negative, statistically significant, but small in magnitude in both the DI and SSI specifications.[21] In both the DI and SSI specifications, our estimated disability program

21. The people that are not employed but not disabled are most likely unemployed and receiving unemployment insurance payments. Thus, while inconsistent with a leisure complementarity story, our results are very consistent with findings in recent papers showing substitutability between social safety net programs (Borghans, Gielen, and Luttmer 2014).

network coefficients remain positive, statistically significant, and of roughly the same magnitude when the interactions are added to the models suggesting that leisure complementarities are not the driving force behind ethnic network effects.

VII. ANALYSIS OF RETIREMENT-AGE SAMPLE

As discussed in the Introduction, we would ideally consider the effect of exposure to disability program participation on take-up among only those who are eligible for the programs. Unfortunately, it is not possible using Census data to perfectly establish eligibility given the lack of information on whether immigrants in our sample are undocumented, whether they have sufficient work histories in Social Security-covered jobs to qualify for DI, and whether they misreport their incomes and assets to the SSA in order to qualify for SSI. Thus, in addition to network effects in take-up conditional on eligibility, our estimated coefficients on the interaction between contact availability and average co-ethnic disability program participation may also measure similarities in eligibility among co-ethnics that live in the same area. For example, even if social norms and information sharing played no role in people's disability program participation, it might be possible to estimate a positive and statistically significant network coefficient in our DI interaction variable if, conditional on country of origin, immigrants who satisfy the work experience requirements of the DI program tend to reside near each other. A similar story can be told for SSI.

Existing analyses do point to similarities in work experience and poverty rates among immigrants from the same country that live surrounded by co-ethnics. There is a large literature documenting how personal connections aid in finding jobs (see Bayer, Ross, and Topa 2008 and references therein). A parallel literature presents evidence of networks in welfare take-up (Åslund and Fredriksson 2009; Bertrand, Luttmer, and Mullainathan 2000) while the results in Brügger, Lalive, and Zweimüller (2012) point to the importance of culture in determining unemployment rates. Given that welfare recipients and the long-term unemployed are less likely to have the work experience necessary to qualify for DI and more likely to satisfy the income and asset constraints for SSI, our estimated network effects may simply reflect the role of social interactions in determining who qualifies for the disability programs as opposed to who participates, conditional on qualifying.

To examine how much of our DI estimated network effects are likely to be explained by work experience, we exploit the fact that the disability and retirement programs of the SSA have almost identical nondisability-related eligibility requirements. In fact, both are part of the same federal program, Old-Age, Survivors, and Disability Insurance. To qualify for Social Security retirement income, individuals above retirement age must satisfy the same work history requirements as DI recipients (they need not satisfy the recent work requirement) but receive benefits, irrespective of disability. Given the magnitude of the program, it is unlikely that there are any significant taboos against receiving retirement income. Moreover, because no evidence of disability is required to receive these benefits, the application process is significantly more straightforward. Thus, a positive and statistically significant interaction coefficient estimate in a model with the receipt of Social Security retirement income as the dependent variable (and average retirement take-up within country of origin in the measure of networks) might be interpreted as evidence that similarities in work histories are driving at least part of our estimated network effects in the working-age sample. Note that the variation in the average take-up across ethnic groups, among non-workers, is likely to come solely from cross-group differences in satisfying the work history requirement.

Similarly, SSI is available to individuals age 65 and above, regardless of disability status, as long as applicants meet the income and asset requirements. Positive and statistically significant estimated coefficients on our interaction variables in models run on this older sample might be suggestive of cultures of poverty which make people eligible for SSI for reasons unrelated to disability.

The Census reports all income received from Social Security during the previous year. As discussed above, this includes pensions, survivors' benefits, permanent disability insurance, and U.S. government Railroad Retirement insurance payments. Our baseline models are restricted to non-widowed immigrants under the age of 62, and so income from Social Security is most likely to be DI income. Our retirement-age sample consists of individuals above the age of 65 and so the same variable measures the receipt of retirement income. Analogously, SSI recipients in the working-age sample receive SSI as a result

TABLE 9

Information Sharing, Retirement-Age Sample, Social Security, and SSI

	Social Security		SSI	
	(1)	(2)	(3)	(4)
CA * proportion of co-ethnics receiving Social Security	0.033** (0.008)	0.010 (0.014)		
CA * proportion of co-ethnics receiving SSI			0.131** (0.012)	0.115** (0.022)
CA	−0.031** (0.005)	−0.012 (0.010)	−0.010** (0.002)	−0.013** (0.003)
CA * proportion of co-ethnics receiving Social Security * HS dropout		0.016 (0.017)		
CA * proportion of co-ethnics receiving Social Security * HS degree		0.030*** (0.017)		
CA * proportion of co-ethnics receiving Social Security * some college		0.042* (0.021)		
CA * proportion of co-ethnics receiving SSI * HS dropout				0.019 (0.029)
CA * proportion of co-ethnics receiving SSI * HS degree				0.007 (0.024)
CA * proportion of co-ethnics receiving SSI * some college				0.041*** (0.025)
Observations	100,090	100,090	100,090	100,090
Adjusted R^2	0.241	0.242	0.162	0.163
Country of origin fixed effects	Yes	Yes	Yes	Yes
PUMA fixed effects	Yes	Yes	Yes	Yes
Age fixed effects	Yes	Yes	Yes	Yes
Years in the U.S. fixed effects	Yes	Yes	Yes	Yes
Controls	Yes	Yes	Yes	Yes

Notes: Coefficients in this table are estimated on a sample of individuals age 65 and above using the same data restrictions as for the baseline sample. Coefficients are estimated using linear probability models. In columns 1 and 2, the dependent variable is receiving Social Security and in columns 3 and 4 the dependent variable is receiving Supplemental Security Income (SSI). Heteroskedasticity corrected standard errors clustered by geography and country of origin cells are in parentheses. The triple interaction coefficients in column 2 (4) are not jointly statistically significant different from zero: F-test = 1.79, p value = .146 (F test = 1.00, p value = .393). For reasons of brevity, we only report coefficients on the variables of interest. Other controls include all the necessary double interactions (for columns 2 and 4 only) as well as the controls shown in Table 2. Estimates are weighted using the appropriate person-level weights provided by the 2000 U.S. Census.

*Significant at 10%; ** significant at 5%; ***significant at 1%.

of a disability but in the over 65 sample, they need not have a disability.

Table 9 shows results of our models run on an age 65 plus sample of non-workers. We examine only non-workers to increase the likelihood that everyone in this sample would receive benefits if eligible; workers may be eligible but choose to postpone applying for Social Security in order to increase future Social Security benefits. As can be seen in columns 1 and 3, estimated coefficients on the interaction term are positive and significant in both the Social Security and the SSI models, but the retirement-age Social Security coefficient is less than a third the size of the network coefficient in the DI model while the retirement-age SSI network coefficient is about half the size of the comparable coefficient in the baseline SSI model. These retirement-age estimates are even smaller when considered relative to the means of the dependent variables. Social security take-up in our retirement sample is 69.3%, which is significantly above DI take-up of 1.7% in our baseline sample. The difference is not as stark for SSI, 1.3% of the baseline sample receives SSI income while only 10.5% of the retirement-age sample receives SSI. We conclude that although similarities in eligibility for the two disability programs seem to explain part of the estimated network effects in our baseline models, they cannot explain the total effect.

Readers may be concerned that the retired sample estimates are underestimating the true impact of eligibility in our working age sample. After all, the retired sample in the 2000 Census consists of a completely different cohort than the working age sample from the same Census. Eligibility may simply be less important for this older cohort. To examine this issue, we computed network effects for a sample of 57- to 61-year-olds using the 2000 Census data and compared those results to network effects computed using data on the same age cohort in the 2008–2010 American Community Surveys (ACS). These two samples reflect essentially the same cohort measured at two points in time: once just before they are eligible for retirement and once shortly after. As can be seen in Table S11 (column 2), the estimated network coefficient for DI in the 57- to 61-year-old Census sample is 0.097 with a p value of .409. In the retirement-age ACS sample, the estimated DI network coefficient is 0.026 with a p value of .441 (column 3). Neither coefficient is statistically significant, potentially because of the relatively small sample sizes, but the magnitude drops substantially just after retirement age. In the SSI models (columns 4-6), the network coefficient dropped from 0.611 (column 5) with a p value smaller than .001, to .231 with a p value smaller than .001 (column 6). Therefore we conclude that the drop in the estimated network coefficients at retirement age cannot be explained by differences across cohorts.

A potential concern with even these estimates, at least in the DI context, is that older immigrants are more likely, all else equal, to have lived in the United States for more years and are therefore more likely to have worked enough years to qualify for Social Security benefits, both retirement benefits and disability benefits. Although we control for years in the U.S. fixed effects in all specifications, this may be problematic if immigrants who have been in the United States for more years are less sensitive to peer effects in becoming eligible for the programs. To examine whether this causes our retirement sample to underestimate eligibility effects, we consider whether our estimated network effects differ with years in the United States in the baseline sample. We find that the estimated coefficient on a triple interaction between contact availability, proportion of co-ethnics receiving DI payments, and years in the United States is practically zero and statistically insignificant suggesting that this should not be much of a concern. The same is true in the SSI model. Results are shown in columns 1 and 4 of Table S11.

As an additional test of whether our estimated network effects in the working age sample are measuring social norms, we re-estimate our WVS models on the retirement-age sample. While cheating the government taboos are likely to play a large role in determining who exaggerates disabilities, they are unlikely to determine who receives Social Security retirement benefits or SSI for people age 65 and above. Presumably, work norms are less important for retirement-age individuals. Using the retirement-age sample, we test whether network effects are in fact weaker for immigrants from countries with stronger taboos against cheating the government and importance of work norms.

Results for Social Security retirement, provided in columns 5 through 8 of Table 6, show that the triple interaction variables have no systematic effect, and all have much smaller magnitudes than the corresponding coefficients constructed using the working-age sample. Results from the SSI model, provided in columns 5 through 8 of Table 7, show that the estimated government cheating taboo triple interaction coefficient is statistically significant at the 10% level. However, in all specifications, magnitudes of coefficients are substantially smaller in the elderly sample than in the working-age sample.[22] We might therefore conclude that norms have a much stronger impact on network effects in disability program participation conditional on eligibility than on determining eligibility.

We also explored whether the relationships we found in the working-age samples between education and estimated network effects disappear in the retirement-age sample. Because older immigrants need not prove the existence of a disability, information sharing should be significantly less important in this sample, and so differences in network effects by education level should be much less pronounced. Recall that in the working-age sample, estimated DI network effects did not vary significantly with education while estimated SSI network effects were stronger among immigrants with less education. The retirement-age analysis, presented in columns 2 and 4 of Table 9, shows that while some of the estimated coefficients on the triple interactions are statistically different from zero, they are not jointly statistically different from zero suggesting no relationship between education and network effects in retirement-age DI or SSI take-up. This implies that income and asset restrictions cannot explain the education patterns seen in the working-age sample. We conclude that information sharing is unlikely to be an

22. We also estimated models using triple interactions constructed from the individual WVS variables. As can be seen in Tables S12 and S13, none of the estimated triple interaction coefficients are statistically different from zero at the 5% level. All estimates are significantly smaller in magnitude than those from the baseline sample shown in Tables S9 and S10.

important determinant of DI take-up but may be important for SSI take-up.

All these results certainly point to a role of networks on disability program participation, but it is useful to think about how the coefficients translate into parameters with policy implications. Specifically, we might ask how much networks magnify the effect of changes in policies which would increase the number of people eligible for disability programs and how these multipliers compare to multipliers for other types of social programs. To answer these questions, we start by adopting the functional form used in Bertrand et al. so that clearer comparisons can be made. We also subtract estimates of the network coefficient in the retirement-age models (0.033 for Social Security and 0.132 for SSI) from those in the baseline models (0.091 for DI and 0.271 for SSI) in order to get clear estimates of network effects in disability program take-up.

Following Bertrand et al.'s methodology for computing network effects and making their assumption that the network estimate entirely reflects an endogenous rather than contextual effect, our estimates imply that network effects would amplify the effects of disability program policy changes by as much as 8.8% for DI and 25.7% for SSI. As might be expected, given that it is necessary to have some type of disability in order to qualify for disability payments, our estimated DI multiplier is significantly smaller and the SSI multiplier slightly smaller than Bertrand et al.'s 27%.

VIII. CONCLUSION

We examine the influence of ethnic networks in determining take-up of publically funded disability programs in the United States. Our results suggest that immigrants who reside around co-ethnics are especially likely to participate in the programs if they belong to high take-up groups. Findings are robust to the inclusion of several aggregate variables. The retiree results suggest that part of our estimated network effects reflect cross-group differences in the likelihood of satisfying the nondisability-related requirements of the two disability programs. We also present evidence suggesting that social norms are important drivers of ethnic network effects in disability program take-up.

Census data do not allow us to determine whether norms matter because people deserving of benefits do not apply unless norms make participation acceptable or because many applicants are in fact capable of working and exaggerate disabilities when taboos become more lax. However, our prior work on SSI take-up using NHIS data (Furtado and Theodoropoulos 2013) is suggestive of taboos against applying for benefits, despite having only marginal disabilities, becoming more lax as more people take up benefits.

We view our results as evidence that social interactions affect disability program take-up in general, but our analysis focuses on immigrants. Regardless of how much of our conclusions can be extrapolated to the general population, studying immigrant take-up of disability programs is interesting in its own right given its relevance to immigration policy. We hope our results are intriguing enough to motivate broader studies of network effects in disability program take-up.

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SUPPORTING INFORMATION

Additional Supporting Information may be found in the online version of this article:

Table S1. PUMA vs. MSA, DI

Table S2. PUMA vs. MSA, SSI

Table S3. Alternative specifications of the network variable and of contact availability

Table S4. Logit and probit estimation

Table S5. Descriptive statistics on aggregate variables

Table S6. Differences in differences and ratio of ratios estimates

Table S7. Descriptive statistics of the World Values Survey variables

Table S8. Principal component analysis (PCA)

Table S9. Social norms—DI, individual WVS variables

Table S10. Social norms—SSI individual WVS variables

Table S11. Robustness checks: years in the U.S. and cohort effects

Table S12. Social norms, retirement-age sample—social security, individual WVS variables

Table S13. Social norms, retirement-age sample—SSI, individual WVS variables