Immigrant Networks and the Take-Up of Disability Programs: Evidence from US Census Data

Delia Furtado, University of Connecticut and Institute for the Study of Labor (IZA)[†]
Nikolaos Theodoropoulos, University of Cyprus[‡]

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Abstract

This paper examines the role of ethnic networks in disability program take-up among working-age immigrants in the United States. We find that even when controlling for country of origin and area of residence fixed effects, immigrants residing amidst a large number of co-ethnics are more likely to receive disability payments when their ethnic groups have higher take-up rates. Although this pattern can be partially explained by cross-group differences in satisfying the work history or income and asset requirements of the disability programs, we also present evidence suggesting that social norms play an important role.

Keywords Social Security Disability Insurance, Supplementary Security Income, Networks, Social norms, Immigrants

JEL Classification: C31, H55, I18, J61

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[†] Delia Furtado, Department of Economics, University of Connecticut, 365 Fairfield Way, Unit 1063, Storrs, CT 06269-1063; E-mail: Delia.Furtado@uconn.edu; Office Phone: 860-486-0615; Fax: 860-486-4463.

[‡] Nikolaos Theodoropoulos, Department of Economics, University of Cyprus, P.O. Box 20537, CY-1678 Nicosia, Cyprus; E-mail: n.theodoropoulos@ucy.ac.cy

1. Introduction

In 2008, 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 135.8 billion dollars in benefits to the disabled (US Census Bureau 2011).[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 2018 if no legislative actions are taken (Congressional Budget Office 2010). 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 strong role in disability program take-up. On the other hand, if the Social Security Administration (SSA) were 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

[1] In comparison, only about 10 billion dollars were paid to Temporary Assistance for Needy Families (TANF) recipients in the same year (US Census Bureau 2011).

SSA were 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 to 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 since 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 identifying the role of networks in the use of publicly-funded prenatal care. Similarly, Deri (2005) and Devillanova (2008) find evidence of networks effects in health care utilization while Gee and Giuntella (2011) find evidence of network effects in the take-up of Medicaid. Aslund and Fredriksson (2009) estimate

[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 percent of Long Island Railroad's disability applications. They were charged with preparing fraudulent medical assessments for hundreds of retirees (Raushbaum and Secret 2011).

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 Social Security Disability Insurance, 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.

To our knowledge, Rege, Telle, and Votruba (2009a) is the only other study of the role of social interactions in disability program participation. Using neighbors' exposure to plant downsizing as an instrument for neighbors' 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. Not only does our paper differ from theirs in terms of empirical approach, but our focus is on immigrant networks within a US context, and we examine two types of disability programs aimed at different populations.

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 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 US. These variables have no impact on our estimated network effects.

The next step in our analysis is to explore how ethnic networks operate. Using a variety of questions from the World Values Survey to measure home country norms, 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. We also present some suggestive evidence that information sharing is 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 amidst 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 non-disability 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 likely to be, 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 non-disability related requirements for 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 show that home country social norms measured in the World Values Survey have small and statistically insignificant impacts on network effects estimated using 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 2 provides background information on the DI and SSI disability programs. Section 3 explains our identification strategy. Section 4 presents the data and Section 5 outlines the main results and addresses concerns about omitted variable bias and selective migration. Sections 6 and 7 examine the mechanisms through which networks operate and conclusions are provided in Section 8.

2. Background on Disability Programs in the US

The Social Security Disability Insurance program was established in 1956 to insure US 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 five of the past ten 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 Supplemental Security Income 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 that the individual has not engaged in substantial gainful activity (SGA), defined in the year 2010 as earning $1000 per month, in the previous five 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 2011). SSI applications have lower approval rates than DI applications (Annual Statistical Report on the Social Security Disability Insurance 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 high earners. DI recipients are also eligible for Medicare coverage after two 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 non-citizens, including those legally in the US. Initially, practically all non-citizens were barred from receiving SSI, but later reforms restored SSI disability benefits to those who were legally in the US on August 22, 1996. All of the immigrants in our sample were residing in the US five years prior to the 2000 Census, and so, as long as they satisfy the other

[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, Buchinsky, Chan, Rust and Sheidvasser (1999).

program requirements and are legally residing in the US, they are eligible for both types of disability programs.[4]

3. Empirical approach

Estimating social interaction effects is difficult both because information on people's social contacts is typically unavailable and because 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 amidst 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 DDS offices and so regional variation in the leniency of DDS 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 in the US with little knowledge of US customs, institutions,

[4] Immigrants arriving in the US 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 seven years in the US. Other non-citizens cannot receive SSI.

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

and language, making it significantly 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 percent 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 amidst 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:

$$D_{ijk} = \beta_1 \bar{D}_j \times CA_{jk} + \beta_2 CA_{jk} + \mathbf{X}_{ijk}\beta_3 + \delta_j + \gamma_k + \varepsilon_{ijk}, \tag{1}$$

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 $\bar{D}_j$.[7] This will refer to average DI take-up in DI models and

[6] PUMAs are the smallest level of geography available in the 5 percent 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 A4A and A4B in the Appendix, results were similar. Dropping PUMAs in the five largest MSAs also yielded similar results (see columns 3 and 4 of Tables A4A and A4B).

[7] Another approach often used in the literature is to construct this average separately by PUMA. That might be a

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

$$\log\left(\frac{C_{jk}}{P_k}\right),$$

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 of the omitted factors that are typically problematic in this type of study. 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 amidst a large number of others with their ethnic background may be very similar to them in ways which 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 on our interaction may simply reflect correlated effects which are a result of unobserved characteristics that affect individuals in

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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.

[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. Our results are robust to dropping the log.

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 running our models on retirement age individuals.

4. Data

Our source of data is the 5 percent sample of the 2000 US 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 the 1996 Welfare Reform Act, we limit our analysis to those immigrants who were living in the US in 1995, five 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 non-citizens are considered immigrants meaning that Puerto Ricans and people from other US 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 as well as individuals whose countries of origin are not clearly specified in the data. Finally, we merged Azores with Portugal, Korea with South Korea, and all of the countries

[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 6 and 7.

comprising the United Kingdom together with each other in order to match the Census data with data from the World Values Survey for our mechanisms analysis.

The US 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 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 (Annual Statistical Report on the Social Security Disability Insurance Program, Chart 12, 2010). We remind readers that the foreign born are significantly less likely to satisfy the DI work history requirements both because they may not have resided in the US for a sufficient number of years and because they are more likely to work “under the table” or not work at all in the years they have resided in the US. Given their typically lower earnings than natives (Larsen 2004), immigrants are also more likely to qualify for SSI. For

[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.

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 US 18.6 years on average, making them very likely to be eligible for DI. In fact, 80 percent of our sample has lived in the U.S. for more than ten 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 US 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 SSI—12.1 percent of DI recipients receive SSI and 15.4 percent of SSI recipients receive DI.

The contact availability variable suggests that on average immigrants in our sample live in PUMAs where 7 percent of the population shares their country of origin. Further examination reveals considerable variation in this variable. About 25 percent of the immigrants in our sample live in PUMAs where less than 0.3 percent of the population is from their country of origin while a little over 5 percent live in PUMAs where more than 30 percent of the population shares their ethnic origin.

5. 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] Our parameters of interest are identified from variation across 130 countries of origin and 2071 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 very low rates of take-up. Estimated coefficients on all of the other control variables have the expected signs. Given that males are more likely to have substantial work histories, it should not be surprising that they are more likely than females to receive DI but less likely to receive SSI. Blacks are more likely than other racial groups to receive disability payments of both types. Hispanics are less likely than whites to receive SSI, but are marginally more likely than whites to receive DI.

In the second and fourth columns, additional demographic and human capital controls--including schooling, years in the US fixed effects and whether the person has a work-preventing disability--are added. Naturally, people with disabilities are more likely to receive disability

[11] We also estimated this equation using probit and logit models without PUMA fixed effects. As can be seen in Appendix Table A5, marginal effects of our network measure were positive and significant at the one percent level.

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 ten percent increase in the proportion of co-ethnics increases the likelihood of going on DI by a statistically indistinguishable from zero 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 ten percent 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 amidst 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 (.013), a ten percent increase in the proportion of co-ethnics actually results in a statistically insignificant .005 decrease in the probability of SSI take-up. However, that same increase in exposure to co-ethnics leads to a statistically significant .211 increase in the probability of take-up for those in groups with the highest SSI take-up, 0.093.

[12] Admittedly, this is a rather crude measure of disability. However, a similar study of SSI take-up using data from the National Health Interview Survey (NHIS) suggests that estimated network effects are robust to models controlling for health behaviors, such as smoking, and more subjective measures of general health (Furtado and Theodoropoulos 2013).

[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).

Our finding that network effects are so much stronger for SSI take-up than DI take-up should not be surprising 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. Presumably, any taboos against exaggerated disability claims should be more important for SSI than DI.

B. Robustness Checks

As discussed above, the main potential threat to our identification strategy is the possibility that immigrants who choose to reside amidst a large number of co-nationals are more similar to the people in their ethnic groups in unobservable ways which then result in similarities in disability program participation. A particular concern is that immigrants residing amidst 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 five years. The disabled typically are 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 simply control for people's listed occupations and industries.[14] However, we do construct several aggregate variables which can be used to alleviate the most obvious occupation-related concerns with our identification strategy.

[14] We did run regressions with occupation fixed effects for both DI and SSI. Estimated network coefficients were the 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 percent of SSI recipients list an occupation in the Census; 66 percent of DI recipients list an occupation.

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 participation. Plant downsizing in Norway has also been found to substantially increase disability program participation of workers in affected plants (Rege, Telle and Votruba 2009b). 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 have 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 (IIF) 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 change our estimated network coefficients. Descriptive statistics on all of these aggregate variables are shown in Appendix Table A1.

[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.

Tables 4A and 4B 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, the addition of 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 of the individual-level variables we control for in our baseline regressions (again, see Appendix Table A1 for descriptive statistics). In the third column of Tables 4A and 4B, we add average values of age, years in the US, and years of schooling 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 of the aggregate variables, and results remain the same. We conclude from these analyses that estimated network effects in these models are very robust.

As a final set of robustness checks (results available upon request), we explored whether our estimated effects are larger for people we would expect to be more socially connected to their groups. We found that immigrants fluent in English are less sensitive to ethnic networks when it

[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.

comes to both DI and SSI participation. The language ability differential is stronger for SSI than DI which makes sense in that poor English speakers without work experience living at or near the poverty level should be especially dependent on information obtained from their ethnic communities. We also compared network effects for the foreign and native born using ancestry instead of country of birth to measure ethnic origin. In both DI and SSI models, results pointed to strong network effects for the foreign born but small effects for the native born which were statistically insignificant in the DI model.

Reverse causality is a potential concern if disability income recipients from ethnic groups with high rates of take-up are especially likely to move to, or less likely to move out of, neighborhoods with large co-ethnic representations after becoming disabled. Unfortunately, the Census contains only limited information on migration patterns and no information on when people applied for and started receiving disability benefits. Nevertheless, we did explore whether immigrants receiving disability payments are more likely to have moved from states with strict DDS offices to states with more lenient DDS offices in order to receive benefits. Our results, shown and discussed in Furtado and Theodoropoulos (2014) do not provide any evidence that this is the case. We do not place much emphasis on these findings, however, given that much of selective migration may have taken more than five years prior to the survey.

6. 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' 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 four-wave integrated data file (WVS) (European and World Values Surveys 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 the 1995 responses. In the end, we were able to match all of our WVS variables of interest to 37 of the 130 countries in our Census sample.[17]

[17] Our sample size falls by 42 percent when limited to observations with non-missing data on all of our WVS

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 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 Appendix Table A2.

Because there are several World Values Survey questions essentially measuring the same concepts, we aggregated information using principal components analysis. As can be seen in Appendix Table A3, 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.

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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 6A). However, the magnitude is roughly the same 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 Table 6B). 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.

Moreover, the first principal component explains more than half of the common variance of the three measures of government cheating taboos and over 80 percent of the common variance of the five measures of work norms. All factors load positively on both of the first principal components. For all of 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:

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

where $\overline{Norm}_j$ 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 5A and 5B present results for DI and SSI respectively. 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 cheating the government taboos and importance of work norms both 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

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

[19] We were also interested in whether exposure to co-ethnics directly led 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.

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. We suspect this may be a result of multicollinearity issues. These estimates suggest that norms do play some 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 -.08 on the importance of work norms in the DI model. Take two immigrants from ethnic groups with the same average DI take-up (.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 ten percent 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 25^{th} 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 75^{th} percentile.

We also ran similar models using the individual variables constructed from the World Values Survey. Findings, shown in Appendix Tables A6A and A6B, 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 programs, how to complete the necessary paperwork, the doctors that provide the most convincing cases for disability, and the lawyers that 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; Aslund and Fredriksson 2009). Unfortunately, we do not have a clean way to test for information sharing with 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 6 presents results from models including interactions of our network measure and educational attainment. Columns 1 and 4 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 than college graduates to exposure to SSI take-up within their ethnic groups, 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 education level, contact availability, and average DI take-up are statistically different from zero, and we cannot reject the null hypothesis that the coefficients are equal to each other. This might 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 5 of Table 6, 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.[20] In both the DI and SSI specifications, our estimated disability program 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.

For further examination of the mechanisms driving our network effects, we explore the sources of cross-ethnicity variation in disability program take-up. For example, take two identical individuals with the same exposure to friends receiving disability payments. However,

[20] 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 2010).

person 1's friends are receiving payments because of observable characteristics (they are older, for example) while person 2's friends are receiving payments for unexplainable reasons (they exaggerate disabilities). We would expect person 2 to be more sensitive to exposure to disability program users than person 1 if social norms are the driving force behind network effects. If on the other hand, network effects were driven solely by information sharing and leisure complementarities, then both person 1 and person 2 should have equal likelihoods of disability program participation.

We examine this by first estimating models of DI and SSI take-up as a function of all of the variables in the baseline model except for the network interaction. We then construct the country of origin-level means of the predicted values of the dependent variables as well as residuals from these models. Finally, we replace the network variable in our baseline model with an interaction between contact availability and average predicted disability and an interaction between contact availability and the average residual variable. Results, shown in columns 3 and 6 of Table 6, suggest that people are in fact very responsive to the unexplained differences in disability take-up by ethnic group and not at all responsive to the explained differences. We view this as further evidence that cultural norms are at least partially driving our results in the baseline models.

7. 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 that 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 Social Security Administration 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 on our DI interaction variable if, conditional on country of origin, immigrants that 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 (Bertrand et al., Aslund and Fredriksson 2009) while the results in Brügger, Lalive and Zweimüller (2009) 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 Social

Security Administration have almost identical eligibility requirements. In fact, both are part of the same federal program, Old-Age, Survivors, and Disability Insurance (OASDI). 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 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.

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

Table 7 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 percent which is significantly above DI take-up of 1.7 in our baseline sample. The difference is not as stark for SSI, 1.3 percent of the baseline sample receives SSI income while only 10.5 percent 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 to 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 Appendix Table A7 (column 2), the estimated network coefficient for DI in the 57-61 year old Census sample is 0.097 with a p-value of 0.409. In the retirement-age ACS sample, the estimated DI network coefficient is 0.026 with a p-value of 0.441 (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 to 6), the network coefficient dropped from 0.611 (column 5) with a p-value of 0.000 to 0.231 with a p-value of 0.000 (column 6). We conclude therefore 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 US 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 US fixed effects in all specifications, this may be problematic if immigrants who have been in the US 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 US 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 US 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 Appendix Table A7.

As an additional test of whether our estimated network effects in the working age sample are measuring social norms, we re-estimate our World Values Survey 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. Importance of work norms are also presumably 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 5A, show that neither of the estimated triple interaction coefficients are statistically different from zero, 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 5B, show that the estimated government cheating taboo triple interaction coefficient is statistically significant at the 10 percent level. However, in all specifications, magnitudes of coefficients are substantially smaller in the elderly sample than in the working-age sample.[21] We might conclude therefore 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

[21] We also estimated models using triple interactions constructed from the individual WVS variables. As can be seen in Appendix Tables A8A and A8B, none of the estimated triple interaction coefficients are statistically different from zero. All estimates are significantly smaller in magnitude than those from the baseline sample shown in Appendix Tables A6A and A6B.

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

All of 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[22] 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 cleaner estimates of network effects in disability program take-up.

Following Bertrand et al's methodology for computing network effects, our estimates imply that network effects would amplify the effects of disability program policy changes by as much

[22] Before taking the log, Bertrand et al. 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.

as 8.8 percent for DI and 25.7 percent 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 the Bertrand et al.'s 27 percent.

8. Conclusion

This paper examines 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 a wide range of control variables. Our analysis of retirees suggests that part of our estimated network effects reflect cross-group differences in the likelihood of satisfying the non-disability related requirements of the two disability programs. However, 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. A related paper (Furtado and Theodoropoulos 2013) finds 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 may suggest that taboos against applying for benefits despite having only marginal disabilities become more lax as more people take-up benefits. However, it does not appear that social networks promote egregious lies on SSI applications that do not get caught during application and appeals process, nor do they operate via information sharing on the most effective appeals lawyers.

We view our results as suggestive of how 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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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 (Seylon)	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	147

Notes: Our sample consists of non-widowed, non-institutionalized immigrants, age 25 to 61, who are not currently in school and who were living in the US five 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 US Census.

Table 2 Descriptive Statistics

	Whole Sample		DI Sample	SSI Sample
	(1)	(2)	(3)	(4)
	Mean	Std. Dev.	Mean	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 US	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 US Census.

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**
	(0.000)	(0.000)	(0.000)	(0.000)
Male	0.001**	0.0004	-0.002**	-0.003**
	(0.000)	(0.000)	(0.000)	(0.000)
Hispanic	0.001+	0.001+	-0.001+	-0.001*
	(0.001)	(0.001)	(0.000)	(0.000)
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**
		(0.000)		(0.000)
Some college		0.003**		0.003**
		(0.001)		(0.000)
Spouse present		-0.008**		-0.013**
		(0.000)		(0.000)
Child		0.001+		-0.001*
		(0.001)		(0.001)
Number of children		-0.001*		0.0002
		(0.000)		(0.000)
English fluency		-0.003**		-0.002**
		(0.000)		(0.000)
Disability		0.013**		0.020**
		(0.001)		(0.001)
Observations	704,871	704,871	704,871	704,871
Adjusted R-squared	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 US 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. Levels of significance: ** significance at 1%, *significance at 5%, + significance at 10%.

Table 4A Effects of Aggregate Variables on 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**
	(0.000)	(0.000)	(0.000)	(0.000)
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
		(0.001)		(0.001)
On-the-job injury rates within PUMA-Country of origin		-0.0004		-0.0005
		(0.000)		(0.000)
On-the-job fatality rates within PUMA-Country of origin		0.0003		0.0003
		(0.000)		(0.000)
Average years of schooling within PUMA-Country of origin			-0.012	-0.011
			(0.013)	(0.014)
Average years in the US 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)
CA * Mean residual DI by country of origin				
Observations	682,378	682,378	682,378	682,378
Adjusted R-squared	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 US 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 number of observations in columns 1 to 4 is less than in column 5 due to missing observations of the aggregate variables. The aggregate injury, fatality, schooling, years in the US, and age variables have been divided by 100. The variable "Mean residual DI by country of origin" in column 5 is constructed by first regressing DI on all of the controls except country of origin, calculating the residuals and then taking the average of the residuals by country of origin. Estimates are weighted using the appropriate person-level weights provided by the 2000 U.S. Census. Levels of significance: ** significance at 1%, *significance at 5%, + significance at 10%.

Table 4B Effects of Aggregate Variables on 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**
	(0.000)	(0.000)	(0.000)	(0.000)
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		-0.001		0.0004
		(0.000)		(0.001)
On-the-job injury rates in PUMA-Country of origin		-0.001**		-0.002**
		(0.000)		(0.000)
On-the-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 US 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)
CA * Mean residual SSI by country of origin				
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 US 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 number of observations in columns 1 to 4 is less than in column 5 due to missing observations of the aggregate variables. The aggregate injury, fatality, schooling, years in the US, and age variables have been divided by 100. The variable "Mean residual SSI by country of origin" in column 5 is constructed by first regressing SSI on all of the controls except country of origin, calculating the residuals and then taking the average of the residuals by country of origin. Estimates are weighted using the appropriate person-level weights provided by the 2000 U.S. Census. Levels of significance: ** significance at 1%, *significance at 5%, + significance at 10%.

Table 5A 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	0.137	0.053	0.041	0.029*	0.039**	0.046**	0.045**
	(0.080)	(0.087)	(0.070)	(0.078)	(0.013)	(0.014)	(0.014)	(0.014)
CA	-0.001	-0.002	-0.001	-0.0004	-0.027**	-0.034**	-0.039**	-0.038**
	(0.001)	(0.002)	(0.001)	(0.001)	(0.008)	(0.009)	(0.009)	(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+		-0.076		-0.013		0.004
		(0.048)		(0.059)		(0.008)		(0.009)
CA * PCA Cheating government taboos		0.001+		0.001		0.006		-0.001
		(0.0006)		(0.001)		(0.004)		(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**	-0.074*			-0.008	-0.008
			(0.028)	(0.032)			(0.007)	(0.007)
CA * PCA work norms			0.001*	0.001+			0.009+	0.010+
			(0.0003)	(0.0004)			(0.005)	(0.005)
Observations	403,288	403,288	403,288	403,288	50,313	50,313	50,313	50,313
Adjusted R-squared	0.015	0.015	0.015	0.015	0.302	0.302	0.302	0.302
Years in the US 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 to 4 use the baseline sample (25 to 61 years old) and columns 5 through to 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 to 4 (5 through to 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 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 column 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. Levels of significance: ** significance at 1%, *significance at 5%, + significance at 10%.

Table 5B 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**	0.266**	0.213**	0.176*	0.140**	0.135**	0.147**	0.146**
	(0.032)	(0.036)	(0.034)	(0.080)	(0.017)	(0.017)	(0.018)	(0.018)
CA	-0.003**	-0.003**	-0.002**	-0.002*	-0.012**	-0.011**	-0.012**	-0.012**
	(0.0004)	(0.0004)	(0.0004)	(0.001)	(0.003)	(0.003)	(0.003)	(0.003)
CA * Proportion of co-ethnics receiving SSI * PCA Cheating government taboos		-0.032*		0.007		-0.013+		0.0004
		(0.014)		(0.035)		(0.008)		(0.009)
CA * PCA Cheating government taboos		0.001**		0.0002		0.003+		-0.001
		(0.0001)		(0.0004)		(0.002)		(0.002)
CA * Proportion of co-ethnics receiving SSI * PCA Work norms			-0.046*	-0.074			-0.007	-0.005
			(0.018)	(0.046)			(0.011)	(0.012)
CA * PCA Work norms			0.0003+	0.0001			-0.003	-0.003+
			(0.0002)	(0.0003)			(0.001)	(0.001)
Observations	403,288	403,288	403,288	403,288	50,313	50,313	50,313	50,313
Adjusted R-squared	0.302	0.302	0.302	0.302	0.222	0.222	0.222	0.222
Years in the US 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 to 4 refer to the baseline sample (25 to 61 years old) and columns 5 through to 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 to 4 (5 through to 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 column 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. Levels of significance: ** significance at 1%, *significance at 5%, + significance at 10%.

Table 6 Information Sharing and Leisure Complementarities, DI and SSI

	DI				SSI	
	(1)	(2)	(3)	(4)	(5)	(6)
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	-0.0002	-0.001**	0.002**	-0.0007
	(0.0004)	(0.001)	(0.0006)	(0.0003)	(0.001)	(0.0009)
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)	
CA * Average Predicted DI/SSI			0.023			0.048
			(0.037)			(0.070)
CA * Average Residual DI/SSI			0.239**			0.383**
			(0.057)			(0.036)
Observations	704,871	704,871	704,871	704,871	704,871	704,871
Adjusted R-squared	0.017	0.017	0.017	0.032	0.029	0.029
Country of origin fixed effects	Yes	Yes	Yes	Yes	Yes	Yes
PUMA fixed effects	Yes	Yes	Yes	Yes	Yes	Yes
Age fixed effects	Yes	Yes	Yes	Yes	Yes	Yes
Years in the US fixed effects	Yes	Yes	Yes	Yes	Yes	Yes
Controls	Yes	Yes	Yes	Yes	Yes	Yes

Notes: In columns 1 and 3 the dependent variable is Disability Insurance (DI) and in columns 4 and 6 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 = 0.841), whereas they are jointly statistically different from zero in Column 4 (F statistic= 12.00, p-value = 0.000). 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 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. Levels of significance: ** significance at 1%, *significance at 5%, + significance at 10%.

Table 7 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**	0.010		
	(0.008)	(0.014)		
CA * Proportion of co-ethnics receiving SSI			0.131**	0.115**
			(0.012)	(0.022)
CA	-0.031**	-0.012	-0.010**	-0.013**
	(0.005)	(0.010)	(0.002)	(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-squared	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 US 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 1 (3) are not jointly statistically significant different from zero: F-test= 1.79, p-value = 0.146 (F-test= 1.00, p-value = 0.393). 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. Levels of significance: ** significance at 1%, *significance at 5%, + significance at 10%.

Appendix

Table A1 Descriptive Statistics on Aggregate Variables

Variable	Mean	Std. Dev.	Min	Max
On-the-job injury rates within PUMA-Country of origin	131.25	64.12	1.66213	1858.876
On-the-job fatality rates within PUMA-Country of origin	7.856	41.35	0	5524.193
Unemployment rate within PUMA-Country of origin	0.058	0.073	0	1
Average years of schooling within PUMA-Country of origin	11.48	3.263	0	22
Average years in the US within PUMA-Country of origin	18.52	5.784	0	61
Average age within PUMA-Country of origin	40.93	4.441	25	61
Yearly wage income within PUMA-Country of origin	32924.83	22463.03	100	385000
Observations		682,378		

Notes: The above descriptive statistics are computed using the 682,378 observations with non-missing data on all of the above variables. The on-the-job fatality rate was computed by dividing the number of fatalities by total employment and then multiplying this figure by 100,000.

Table A2 Descriptive Statistics of the World Values Survey Variables

	Mean	Std. Dev.	Min	Max
Proportion of home country believing that…				
Cheating on taxes never justifiable	0.677	0.119	0.258	0.985
Claiming government benefits to which not entitled never justifiable	0.513	0.138	0.280	0.921
Avoiding fare on public transport never justifiable	0.552	0.161	0.211	0.964
Work is very important in life	0.778	0.150	0.435	0.958
Work is what makes life worth living	0.441	0.137	0	0.781
Proportion of home country strongly agreeing that...				
People who do not work turn lazy	0.335	0.073	0.038	0.660
To develop talents you need to have a job	0.308	0.087	0.148	0.779
Work is a duty towards society	0.279	0.078	0.086	0.764
Work should come first even if it means less spare time	0.252	0.069	0.030	0.694
Observations		403,288		

Notes: The above descriptive statistics are computed using the 403,288 observations with non-missing observations on all of the above variables.

Table A3 Principal Component Analysis (PCA)

	(1)	(2)	(3)	(4)
Panel A. PCA analysis on "Cheating the Government Variables"				
Component	Eigenvalue	Difference	Proportion	Eigenvector (Compoenent1)
Component1	2.416	1.994	0.805	0.555
Component2	0.422	0.261	0.141	0.556
Component3	0.161	.	0.054	0.610
Panel B. PCA analysis on "Importance of Work Variables"				
Component	Eigenvalue	Difference	Proportion	Eigenvector (Component1)
Component1	3.180	1.749	0.530	0.455
Component2	1.432	0.799	0.239	0.367
Component3	0.633	0.288	0.106	0.404
Component4	0.345	0.029	0.058	0.200
Component5	0.316	0.022	0.053	0.454
Component6	0.094	.	0.016	0.499

Notes: Panel A shows the PCA analysis for "Cheating the government" variables. The first principal component is the weighted average of these three variables where the weights are chosen to maximize the proportion of the variance of these three individual variables that can be explained by the first principal component. Column 1 shows the eigenvalue for each component, Column 2 the vertical difference between the components, Column 3 shows the proportion of the variance that can be explained by each component. Finally, Column 4 shows the eigenvector of the first component. Panel B shows the PCA analysis for the "Importance of work" variables. Since we use six variables to proxy importance of work, PCA reports six components.

Table A4A: PUMA vs. MSA, DI

	(1)	(2)	(3)	(4)
			Dropping 5 Largest MSAs	
CA * Proportion of co-ethnics receiving DI	0.097**		0.102**	
	(0.025)		(0.035)	
CA	-0.001**		-0.001*	
	(0.000)		(0.001)	
CA at the MSA level * Proportion of co-ethnics receiving DI		0.072*		0.080*
		(0.028)		(0.039)
CA at the MSA level		-0.001+		-0.001
		(0.001)		(0.001)
Observations	640,930	640,930	407,039	407,039
Adjusted R-squared	0.016	0.015	0.017	0.016
Years in the US fixed effects	Yes	Yes	Yes	Yes
PUMA/MSA fixed effects	Yes	Yes	Yes	Yes
Age fixed effects	Yes	Yes	Yes	Yes
Years in the US fixed effects	Yes	Yes	Yes	Yes
Controls	Yes	Yes	Yes	Yes

Notes: The variable "CA at the MSA level" is constructed at the MSA level as opposed at the PUMA level. Column 1 shows results from PUMA-level regressions run on a sample of immigrants residing in MSAs. Columns 1 and 3 control for PUMA fixed effects, while columns 2 and 4 control for MSA fixed effects. See Table 2 of the paper for the list of controls.

Table A4B: PUMA vs. MSA, SSI

	(1)	(2)	(3)	(4)
			Dropping 5 Largest MSAs	
CA * Proportion of co-ethnics receiving SSI	0.274**		0.297**	
	(0.031)		(0.042)	
CA	-0.004**		-0.004**	
	(0.000)		(0.001)	
CA at the MSA level * Proportion of co-ethnics receiving SSI		0.341**		0.378**
		(0.044)		(0.061)
CA at the MSA level		-0.005**		-0.005**
		(0.001)		(0.001)
Observations	640,930	640,930	407,039	407,039
Adjusted R-squared	0.029	0.027	0.029	0.028
Years in the US fixed effects	Yes	Yes	Yes	Yes
PUMA/MSA fixed effects	Yes	Yes	Yes	Yes
Age fixed effects	Yes	Yes	Yes	Yes
Years in the US fixed effects	Yes	Yes	Yes	Yes
Controls	Yes	Yes	Yes	Yes

Notes: The variable "CA at the MSA level" is constructed at the MSA level as opposed at the PUMA level. Column 1 shows results from PUMA-level regressions run on a sample of immigrants residing in MSAs. Columns 1 and 3 control for PUMA fixed effects, while columns 2 and 4 control for MSA fixed effects. See Table 2 of the paper for the list of controls.

Table A5 Logit and Probit Estimation

	DI				SSI			
	Logit		Probit		Logit		Probit	
	Coef	ME	Coef	ME	Coef	ME	Coef	ME
CA * Proportion of co-ethnics receiving DI /SSI	5.453**	0.087**	2.361**	0.091**	6.510**	0.080**	3.165**	0.093**
	(1.195)	(0.019)	(0.506)	(0.019)	(0.791)	(0.009)	(0.351)	(0.010)
CA	-0.112**	-0.002**	-0.048**	-0.002**	-0.085**	-0.001**	-0.041**	-0.001**
	(0.024)	(0.0003)	(0.506)	(0.0003)	(0.018)	(0.0002)	(0.007)	(0.0002)
Observations	704479		704479		703941		703941	
Log-pseudolikelihood	-1183775.6		-1184446.2		-899194.69		-900239.67	
Country of origin fixed effects	Yes		Yes		Yes		Yes	
PUMA fixed effects	No		No		No		No	
Age fixed effects	Yes		Yes		Yes		Yes	
Years in the US fixed effects	Yes		Yes		Yes		Yes	
Controls	Yes		Yes		Yes		Yes	

Notes: This table shows results from the baseline specification (Columns 2 and 4 of Table 3) estimated using logit and probit models. However, we do not include PUMA fixed in these specifications. In the SSI models, country of origin often predicts SSI perfectly and so observations from certain countries of origin are dropped from the models. Heteroskedasticity corrected standard errors clustered by geography and country of origin cells are in parentheses. Observations are weighted using the appropriate person-level weights provided by the 2000 US Census. Significance levels are noted by the following: ** significance at 1%.

Table A6A Social Norms – DI, Individual WVS Variables

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
CA * Proportion of co-ethnics receiving DI	1.348**	-0.132	0.487+	1.047**	0.291+	0.683**	0.751*	0.631**	0.569**
	(0.407)	(0.347)	(0.291)	(0.349)	(0.172)	(0.263)	(0.302)	(0.244)	(0.204)
CA * Proportion of co-ethnics receiving DI * Cheating on taxes never justifiable	-1.916**								
	(0.576)								
CA * Proportion of co-ethnics receiving DI * Claiming government benefits to which not entitled never justifiable		0.411							
		(0.572)							
CA * Proportion of co-ethnics receiving DI * Avoiding fare on public transport never justifiable			-0.704						
			(0.462)						
CA * Proportion of co-ethnics receiving DI * Work is very important in life				-1.407**					
				(0.474)					
CA * Proportion of co-ethnics receiving DI * Work is what makes life worth living					-0.596				
					(0.448)				
CA * Proportion of co-ethnics receiving DI * People who do not work turn lazy						-1.684*			
						(0.690)			
CA * Proportion of co-ethnics receiving DI * To develop talents you need to have a job							-1.949*		
							(0.821)		
CA * Proportion of co-ethnics receiving DI * Work is a duty towards society								-2.041**	
								(0.789)	
CA * Proportion of co-ethnics receiving DI * Work should come first even if it means less spare time									-1.740**
									(0.670)
Observations	403,288	403,288	403,288	403,288	403,288	403,288	403,288	403,288	403,288
Adjusted R-squared	0.015	0.015	0.015	0.015	0.015	0.015	0.015	0.015	0.015
Clusters	25738	25738	25738	25738	25738	25738	25738	25738	25738
Country of Origin fixed effects	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Years in the US fixed effects	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
PUMA fixed effects	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Age fixed effects	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Controls	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes

Notes: The dependent variable is Disability Insurance (DI) and the estimation method is a linear probability model. All the necessary double interactions are included as well as the controls shown in Table 2 of the paper. Heteroskedasticity corrected standard errors clustered by geography and country of origin cells are in parentheses. Observations are weighted using the appropriate person-level weights provided by the 2000 US Census. Due to missing observations in the WVS data, the coefficients in this table are estimated using 403,288 observations as opposed to 704,871 observations in the baseline sample. Significance levels are noted by the following: ** significance at 1%, * significance at 5%, + significance at 10%.

Table A6B Social Norms – SSI Individual WVS Variables

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
CA * Proportion of co-ethnics receiving SSI	0.470**	0.514*	0.483**	1.121**	0.382**	0.930**	0.146	0.332**	0.563**
	(0.127)	(0.210)	(0.101)	(0.303)	(0.0930)	(0.266)	(0.186)	(0.0767)	(0.173)
CA * Proportion of co-ethnics receiving SSI * Cheating on taxes never justifiable	-0.325+								
	(0.171)								
CA * Proportion of co-ethnics receiving SSI * Claiming government benefits to which not entitled never justifiable		-0.442							
		(0.312)							
CA * Proportion of co-ethnics receiving SSI * Avoiding fare on public transport never justifiable			-0.371**						
			(0.137)						
CA * Proportion of co-ethnics receiving SSI * Work is very important in life				-1.514**					
				(0.489)					
CA * Proportion of co-ethnics receiving SSI * Work is what makes life worth living					-0.448+				
					(0.246)				
CA * Proportion of co-ethnics receiving SSI * People who do not work turn lazy						-2.315*			
						(0.945)			
CA * Proportion of co-ethnics receiving SSI * To develop talents you need to have a job							0.200		
							(0.439)		
CA * Proportion of co-ethnics receiving SSI * Work is a duty towards society								-0.265	
								(0.230)	
CA * Proportion of co-ethnics receiving SSI * Work should come first even if it means less spare time									-1.380+
									(0.712)
Observations	403,288	403,288	403,288	403,288	403,288	403,288	403,288	403,288	403,288
Adjusted R-squared	0.022	0.022	0.022	0.022	0.022	0.022	0.022	0.022	0.022
Country of origin fixed effects	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Years in the US fixed effects	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
PUMA fixed effects	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Age fixed effects									
Controls	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes

Notes: The dependent variable is Supplemental Security Income (SSI) and the estimation method is a linear probability model. All the necessary double interactions are included as well as the controls shown in Table 2 of the paper. Heteroskedasticity corrected standard errors clustered by geography and country of origin cells are in parentheses. Observations are weighted using the appropriate person-level weights provided by the 2000 US Census. Due to missing observations in the WVS data, the coefficients in this table are estimated using 403,288 observations as opposed to 704,871 observations in the baseline sample. Significance levels are noted by the following: ** significance at 1%, * significance at 5%, + significance at 10%.

Table A7 Robustness Checks: Years in the US and Cohort Effects

	DI or Social Security			SSI		
		Age 57-61 in the 2000 cohort			Age 57-61 in the 2000 cohort	
	Baseline sample	Census 2000	ACS 2008-2010 Social Security	Baseline sample	Census 2000	ACS 2008-2010
	(1)	(2)	(3)	(4)	(5)	(6)
CA * Proportion of co-ethnics receiving DI (Social Security in col 3)	0.087*	0.097	0.026			
	(0.043)	(0.118)	(0.034)			
CA * Proportion of co-ethnics receiving DI * Years in the US	-0.0004					
	(0.002)					
CA * Years in the US	0.0001**			0.0001**		
	(0.000)			(0.000)		
Proportion of co-ethnics receiving DI * Years in the US	0.010					
	(0.011)					
CA	-0.003**	0.001	-0.017	-0.005**	-0.008**	-0.008**
	(0.001)	(0.002)	(0.025)	(0.001)	(0.001)	(0.003)
CA * Proportion of co-ethnics receiving SSI				0.309**	0.611**	0.231**
				(0.051)	(0.106)	(0.035)
CA * Proportion of co-ethnics receiving SSI * Years in the US				-0.002		
				(0.002)		
Proportion of co-ethnics receiving SSI * Years in the US				-0.016		
				(0.013)		
Observations	704,871	55,829	23,302	704,871	55,829	23,302
Adjusted R-squared	0.017	0.032	0.273	0.029	0.074	0.211
Country of origin fixed effects	Yes	Yes	Yes	Yes	Yes	Yes
PUMA fixed effects	Yes	Yes	Yes	Yes	Yes	Yes
Age fixed effects	Yes	Yes	Yes	Yes	Yes	Yes
Years in the US FE	Yes	Yes	Yes	Yes	Yes	Yes
Controls	Yes	Yes	Yes	Yes	Yes	Yes

Notes: Columns 1and 4 use the baseline sample, 25-61 year olds, from the 2000 Census. Columns 2 and 5 use data on a cohort aged 57-61 in the year 2000. Columns 3 and 6 use data on the same age cohort in the 2008 to 2010 American Community Survey (ACS). Samples in columns 2, 3, 5 and 6 reflect the same cohort measured at two points in time, once just before they are eligible for retirement and once shortly after. The proportion of co-ethnics receiving SSI is calculated using the working-age sample in columns 4 and 5 but the retirement-age sample in column 6. Other controls include those shown in Table 2 of the paper. 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. Levels of significance: ** significance at 1%, *significance at 5%, + significance at 10%.

Table A8A Social Norms, Retirement Age Sample – Social Security, Individual WVS Variables

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
CA * Proportion of co-ethnics receiving Social Security	0.089	0.101	0.094*	0.095	0.082*	0.061	0.058	0.017	0.092*
	(0.059)	(0.081)	(0.045)	(0.083)	(0.034)	(0.047)	(0.041)	(0.033)	(0.038)
CA * Proportion of co-ethnics receiving Social Security * Cheating on taxes never justifiable	-0.089								
	(0.084)								
CA * Proportion of co-ethnics receiving Social Security * Claiming government benefits to which not entitled never justifiable		-0.106							
		(0.121)							
CA * Proportion of co-ethnics receiving Social Security * Avoiding fare on public transport never justifiable			-0.104						
			(0.069)						
CA * Proportion of co-ethnics receiving Social Security * Work is very important in life				-0.110					
				(0.143)					
CA * Proportion of co-ethnics receiving Social Security * Work is what makes life worth living					-0.121				
					(0.099)				
CA * Proportion of co-ethnics receiving Social Security * People who do not work turn lazy						-0.095			
						(0.165)			
CA * Proportion of co-ethnics receiving Social Security * To develop talents you need to have a job							-0.064		
							(0.115)		
CA * Proportion of co-ethnics receiving Social Security * Work is a duty towards society								0.110	
								(0.122)	
CA * Proportion of co-ethnics receiving Social Security * Work should come first even if it means less spare time									-0.244
									(0.164)
Observations	50,313	50,313	50,313	50,313	50,313	50,313	50,313	50,313	50,313
Adjusted R-squared	0.270	0.270	0.270	0.270	0.270	0.270	0.270	0.270	0.270
Country of origin fixed effects	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Years in the US fixed effects	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
PUMA fixed effects	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Age fixed effects	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Controls	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes

Notes: The dependent variable is Social Security (DI) and the estimation method is a linear probability model. All the necessary double interactions are included as well as the controls shown in Table 2 of the paper. Heteroskedasticity corrected standard errors clustered by geography and country of origin cells are in parentheses. Observations are weighted using the appropriate person-level weights provided by the 2000 US Census. Due to the missing observations in the WVS raw data, the coefficients in this Table are estimated on 50,313 observations as opposed to 100,090 observations in the retirement age sample. Significance levels are noted by the following: ** significance at 1%, * significance at 5%, + significance at 10%.

Table A8B Social Norms, Retirement Age Sample – SSI, Individual WVS Variables

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
CA * Proportion of retirement-age co-ethnics receiving SSI	0.193**	0.256**	0.205**	-0.063	0.177**	0.116	0.079	0.154**	0.210**
	(0.054)	(0.079)	(0.044)	(0.140)	(0.036)	(0.075)	(0.069)	(0.041)	(0.054)
CA * Proportion of retirement-age co-ethnics receiving SSI * Cheating on taxes never justifiable	-0.099								
	(0.083)								
CA * Proportion of retirement-age co-ethnics receiving SSI * Claiming government benefits to which not entitled never justifiable		-0.192							
		(0.130)							
CA * Proportion of retirement-age co-ethnics receiving SSI * Avoiding fare on public transport never justifiable			-0.128+						
			(0.074)						
CA * Proportion of retirement-age co-ethnics receiving SSI * Work is very important in life				0.355					
				(0.240)					
CA * Proportion of retirement-age co-ethnics receiving SSI * Work is what makes life worth living					-0.147				
					(0.129)				
CA * Proportion of retirement-age co-ethnics receiving SSI * People who do not work turn lazy						0.103			
						(0.273)			
CA * Proportion of retirement-age co-ethnics receiving SSI * To develop talents you need to have a job							0.178		
							(0.185)		
CA * Proportion of retirement-age co-ethnics receiving SSI * Work is a duty towards society								-0.009	
								(0.156)	
CA * Proportion of retirement-age co-ethnics receiving SSI * Work should come first even if it means less spare time									-0.344
									(0.255)
Observations	50,313	50,313	50,313	50,313	50,313	50,313	50,313	50,313	50,313
Adjusted R-squared	0.187	0.187	0.187	0.187	0.187	0.187	0.187	0.187	0.187
Country of origin fixed effects	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Years in the US fixed effects	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
PUMA fixed effects	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Age fixed effects	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Controls	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes

Notes: The dependent variable is Supplemental Security Income (SSI) and the estimation method is a linear probability model. All the necessary double interactions are included as well as the controls shown in Table 2 of the paper. For reasons of brevity, we only report the triple interaction coefficients. Heteroskedasticity corrected standard errors clustered by geography and country of origin cells are in parentheses. Observations are weighted using the appropriate person-level weights provided by the 2000 US Census. Due to the missing observations in the WVS raw data, the coefficients in this Table are estimated on 50,313 observations as opposed to 100,090 observations in the retirement age sample. Significance levels are noted by the following: ** significance at 1%, * significance at 5%, + significance at 10%.