DEPARTMENT OF ECONOMICS
UNIVERSITY OF CYPRUS

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Discussion Paper 10-2012

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

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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 find that social norms and, to a lesser extent, information sharing play important roles.

Keywords Social Security Disability Insurance, Supplementary Security Income, Networks, Immigrants
JEL Classification: C31, H55, I18, J61

The research reported here was performed pursuant to a grant from the US Social Security Administration (SSA) funded as part of the Retirement Research Consortium (RRC). The opinions and conclusions expressed are solely those of the authors and do not represent the opinions or policy of SSA, any agency of the federal government, the RRC, the University of Connecticut, the University of Cyprus, or Boston College.

Acknowledgements: For the many valuable comments and suggestions, we would like to thank seminar participants at Lancaster University, Louisiana State University, Northeastern University, University of Connecticut, and Vassar College, as well as conference participants at the Yale University Health Workshop, EALE (Cyprus), PAA meetings in San Francisco, and the NARSC meetings in Miami.

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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 in the past twenty years, the two programs have grown substantially both in terms of benefits per recipient and number of recipients (Autor and Duggan 2006; Bound and Burkhauser 1999; 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 fairly and how any policy changes may ultimately impact disability program take-up.[2] 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 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

[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). Both the DI and SSI programs provide cash benefits to individuals unable to work as a result of a disability and both have the same standards for determining who is disabled, but DI recipients must satisfy certain work history requirements. The SSI program, on the other hand, does not have prior work requirements but does have income and asset limits.

[2] Benitez-Silva, Buchinsky and Rust (2004) estimate that 20 percent of the DI/SSI applicants who receive benefits are not disabled while 60 percent of the applicants who are disabled are denied benefits. See Autor (2011) for a discussion of the major difficulties facing the DI program as well as an analysis of possible policy changes.

the application process is sufficiently complex, then information sharing within social networks may be an important determinant of take-up among the truly disabled.[3] Also, if the Social Security Administration does not screen applicants effectively, 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, which are likely to increase as more people are not working. Moreover, 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 cross-neighborhood differences in labor markets or initial allowance rates by Disability Determination Services (DDS) offices, for example. Another approach to identifying networks might be to examine the relationship between individual outcomes of immigrants and average behaviors in their country of origin groups.[4] While only 1.6 percent of the 25 to 61 year old immigrants in our sample receive DI payments, the proportion ranges from 4.5 among Cape Verdeans to 0.3 among immigrants from New Zealand. The ethnic variation in the proportion receiving SSI is even greater, ranging from 7.3 for Cambodians to practically zero for Norwegians (see Table 1). This also cannot be

[3] Network members may also share information about doctors who are most likely to exaggerate disabilities. According to a recent 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).

[4] For example, Borjas and Lynette (1996) find that benefits received by earlier cohorts of immigrants from an immigrant's country of origin is predictive of the likelihood that a recent arrival will receive a particular type of benefit.

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 types of issues, we use an empirical approach similar to the one pioneered by Bertrand, Luttmer, and Mullainathan (2000) in their study of welfare take-up. Welfare use within language groups is used to measure the group's views and knowledge of welfare programs in the US. To measure the ease with which individuals can be in contact with co-ethnics, the authors use the number of people in a person's local area that speak the same language. The paper's main question is whether being surrounded by people who speak the same language increases welfare use more for people in high welfare-receiving language groups. This approach allows the authors to control for both language group and local area fixed effects which eliminates many of the standard sources of bias.

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) study welfare using a similar estimating equation as Bertrand et al. but exploit the plausibly exogenous placement of refugees in Sweden for tighter identification.

In our analysis, we start by examining whether there are network effects in disability program take-up. 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 both the DI and SSI programs.[5]

Our analysis of Census 2000 data provides evidence of social interactions 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 assimilation and human capital measures 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.

A potential concern 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. To examine how problematic this is likely to be, we construct for each country of origin-local area cell, on-the-job injury rates and unemployment rates. Adding these variables to our baseline models has no impact on our estimated network effects. We also show that estimated effects are stronger for people we would expect to be more socially connected to their ethnic groups—for example, those with worse English speaking abilities.

The next step in our analysis is to explore how ethnic networks may operate. Using data from the World Values Survey, we show that immigrants from countries where people tend to believe that receiving government benefits to which they are not entitled can be justifiable are more likely to receive disability benefits when they reside amidst many co-ethnics. This result certainly points to a potential role of social norms.

To examine the role of information sharing, we estimate models separately by educational attainment under the assumption that people with more education are

[5] As a percentage of gross domestic product (GDP), Norway spends about four times more on disability programs than the US (Social Security Administration 2006), and so any conclusions about disability programs in Norway may not be applicable to the US.

able to navigate application processes without as much need for information gathered from social networks. While our results are consistent with an information sharing story for SSI recipients, there is no clear relationship between education and our estimated network effects for DI recipients.

We also find that, conditional on ethnic group disability program take-up, immigrants in groups with low employment rates are *not* more likely to take-up disability programs when they reside amidst many co-ethnics. Given that leisure is likely to be just as enjoyable if spent with non-disabled co-ethnics that are not working as with non-working disabled co-ethnics, this result suggests that leisure complementarities are not likely to be driving our network effect results.

Our analysis ends with an exploration of whether differences in eligibility for the DI and SSI programs are driving our results. 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 information sharing about the appeals process and social norms about exaggerating a disability do not play any role for these older immigrants, we interpret any estimated network effects 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 substantially smaller but statistically significant estimated network effects in our retirement age sample suggesting that eligibility differences are important but not the sole drivers of our results. We also show that, in contrast to our baseline sample, home country social norms measured in the World Values Survey have no impact on this population. While the clear education patterns in network effects for SSI take-up are not found in the retirement age sample, the DI patterns do appear in the retirement age sample, again suggesting that information sharing may be important for SSI take-up but not DI.

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. Robustness checks are conducted in Section 6, and Section 7 examines the mechanisms through which networks operate. 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.[6] 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

[6] The Social Security Administration measures a quarter of work based on earnings as opposed to time spent working. In the year 2010, workers accumulated one quarter of work experience for every $1120 earned within the year, with a maximum of four quarters which can be earned in any one year. This implies that if a worker were employed the entire year but only earned $1120, that worker would only have accumulated one quarter of experience. On the other hand, if a worker earned $4480 in one month and did not work for the rest of the year, then he will have accumulated the entire four quarters for the year 2010. For the oldest workers, the duration of work requirement generally translates into ten years of work experience while the recent work requirement translates into having worked five of the past ten years.

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 which 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.[7] Roughly one third of all DI applications are awarded initially and about two thirds of all applications are awarded after 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

[7] Rejected applicants can 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, Man Chan, Rust and Sheidvasser (1999).

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 program requirements and are legally residing in the US, they are eligible for both types of disability programs.[8]

3. Empirical approach

An ideal study of the effect of networks on disability program participation would involve randomly assigning some people to a group of friends with high disability program participation and others to a group with low disability program participation. In practice, researchers do not generally have information on people's social circles, much less disability program usage within those circles, and natural experiments which mimic randomly assigning people to groups of friends are difficult to find. 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.

[8] 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 eligible during their first seven years in the US only. Other non-citizens cannot receive SSI.

One often made assumption in the social interactions literature is that people are more likely to befriend those who live geographically close to them. In the context of disability program participation, 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. 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.[9] Also, people living in the same areas are subject to similar levels of pollution and face similar opportunities to purchase unhealthy foods, both of which have detrimental effects on health.[10]

Another problem, given the income and asset limits of SSI and the progressive nature of DI payments, is that people living in the same areas participate in the same labor markets. Workers in areas with few labor market opportunities are more likely to qualify for SSI and are likely to find DI payments more attractive. 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).

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[9] 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-Siva, Buchinsky and Rust 2004). It seems unlikely that these differences are attributable completely to differences in disability rates.

[10] There is a large literature documenting the detrimental effects of pollution on infant health (Currie, Neidell and Schmieder 2009; Currie, Greenstone and Moretti 2011; Currie and Walker 2011). Proximity to fast food restaurants has been shown to positively affect obesity rates (Currie, Della Vigna, Moretti and Vikram 2010).

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, and language, making it significantly easier to interact with others from the same country of origin as opposed to natives or immigrants from different countries. Again, it may be tempting to simply regress disability program participation on the proportion of immigrants from one's country of origin receiving disability payments, but similar problems arise. People from the same country of origin may have similar genetic predispositions to certain health conditions or engage in similar health-related habits related to diet and exercise. Moreover, given any within-ethnicity correlations in occupations, economy-wide shocks to particular industries may have disproportionate effects on certain ethnic groups.

To address these issues, we use an approach pioneered in Bertrand et al.'s (2000) 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. In doing so, we 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:

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

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).[11] The proportion of people receiving disability payments in a person's ethnic group is

[11] 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 (MSA) level, and results were similar.

denoted $\bar{D}_j$.[12] This will refer to average DI take-up in DI models and average SSI take-up in SSI models. CA_{jk} refers to contact availability or the density of country of origin group j in area k. Contact availability is defined as

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

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. 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 controls.

This specification addresses many of the typical concerns associated with this type of analysis. Area of residence fixed effects control for factors related to a person's environment which affect all people living in the same area. Country of origin fixed effects control for all of the unobserved determinants of program take-up which vary by ethnicity. The contact availability variable controls for preferences for living around co-ethnics which might be correlated with disability program take-up in a manner similar across ethnic groups. Our measure of networks, however, 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.

[12] Another approach often used in the literature is to construct this average separately by PUMA. That might be a better measure of disability program take-up among the coethnics 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.

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. Given the restrictions on SSI eligibility imposed by the 1996 Welfare Reform Act, we limit our analysis to those immigrants who were in the US five years prior to responding to the 2000 survey. This restriction also increases the proportion of the sample eligible for DI payments given the program's work history conditions. We keep only immigrants from origin countries with more than 500 observations in the data in order to limit measurement error in our contact availability variable. Only naturalized citizens and non-citizens are considered immigrants. Thus, Puerto Ricans and people from other US territories as well as individuals born abroad of American parents are dropped from the sample.

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 early retirement age. We also drop widows and widowers from the sample to make it less likely that they are receiving survivor benefits.[13] 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 should be

[13] 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.

receiving it as a result of a disability. Our final sample consists of 692,066 observations.

Table 2 shows descriptive statistics of the variables used in the analysis. The proportions of our sample that receive DI and SSI are about equal. 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. Another explanation relates to how benefits are calculated. In order to qualify for DI benefits, the oldest individuals must work approximately ten years, but payments are calculated based on average earnings within the worker's best thirty-five years. Years in which immigrants do not work are counted as zeros. Thus, immigrants with marginal disabilities may choose to forego disability payments, at least until they have worked a substantial number of years in the US.[14] Given their typically lower earnings than natives (Larsen 2004), immigrants are more likely to qualify for SSI. For further details on how immigrants compare to natives in terms of SSI receipt, see Kaushal (2010) which examines elderly immigrants' labor supply responses to changes in SSI requirements in 1996.

Table 2 also shows that on average, disability payment recipients are older, have lower levels of education, and are more likely to live in PUMAs with a large representation of co-ethnics. Immigrants in our sample have lived in the US approximately 19 years, making them very likely to be eligible for DI. Racial

[14] See Gustman and Steinmeier (2000) and, more recently, Borjas (2011) for an examination of how the Social Security benefit formula affects natives and immigrants differently.

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 percent of DI recipients receive SSI and 16 percent of SSI recipients receive DI.

5. Baseline Results

Tables 3A and 3B present estimates of the coefficients in equation (1) for models explaining DI and SSI participation, respectively. Our parameters of interest are identified from variation across 95 countries of origin and 2071 PUMAs. Standard errors are clustered on country of birth-PUMA cells throughout.

As can be seen in the first column of both tables, our estimates 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 which contain only the controls necessary for interpreting the interaction coefficient. In the second column of both tables, basic demographic controls are added to the specification. All estimated coefficients on the controls 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. Married people are less likely to receive both types of disability payments. Blacks are more likely than other racial groups to receive both types of disability payments. Hispanics are less likely than whites to receive SSI, but they do not have statistically different take-up rates of DI. When adding these controls to the basic specification, the estimated interaction coefficient decreases by 14 percent in the

DI model and four percent in the SSI model. The bulk of the decreases are driven by the age fixed effects.

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, such as lower potential wages, which make them significantly more likely to find DI attractive than the New Zealanders who live in New Zealand neighborhoods or other Cape Verdeans who do not live in Cape Verdean neighborhoods.[15] We will devote much of the remaining part of the paper to addressing this type of concern, but as a preliminary check, it is useful to see what happens to our estimated network coefficients when measures of education and assimilation are added to the models. As can be seen in the third column of Tables 3A and 3B, immigrants with more education and better English speaking abilities are less likely to be receiving DI and SSI. Years in the US have a consistently positive effect on the likelihood of receiving DI but a nonlinear effect, increasing in the first 15 years but decreasing thereafter, on the probability of receiving SSI. More importantly, when these variables are added to the model, the network coefficients do not change substantially, in both the DI and SSI specifications. This suggests that the country of origin and PUMA fixed effects are likely to be already controlling for the most influential unobservable characteristics.

These results certainly point to a role of networks on disability program participation, but it is particularly useful to think about how the coefficients translate into parameters with policy implications. If disability program participation results in feedback effects, then we might ask how much networks

[15] Note from Table 1 that Cape Verdeans have the highest DI usage while New Zealanders have the lowest DI usage.

magnify the effect of changes in policies which would increase the number of people eligible for disability programs.[16] As derived in Bertrand et al. (2000), equation (1) implies that a policy which increases disability program participation by one percentage point in a world with no network effects would actually increase participation by $1/(1-\beta_1\overline{CA}_j))$ percentage points for people from country of origin *k*. Taking the weighted mean of this expression over all countries of origin and plugging in our estimates of β_1 from the third columns of Tables 3A and 3B, we conclude that network effects amplify the effects of policy changes by as much as 29 percent for DI and 51 percent for SSI.

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 surely 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.

Despite the fact that it is necessary to have some type of disability in order to qualify for disability payments, our estimated DI multiplier of 1.29 is very similar to the Bertrand et al. (2000) welfare multiplier of 1.27 while our SSI multiplier is larger. We note, however, that we include only the foreign born in our sample while Bertrand et al. examine all people who speak a language other than English at home. Their estimated network coefficients almost double when they focus on a foreign born sample.

[16] According to Manski (1993) the existence of a multiplier rests on the assumption that our estimated network effects are generated from endogenous effects. In Section 7, we will examine the mechanisms driving our results, but for now, readers may interpret our estimated multiplier as an upper bound estimate of the true multiplier.

6. Can Results Be Explained Entirely by Omitted Variable Bias?

A. The Role of Occupational Similarities

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 may resemble their ethnic groups in ways which result in higher disability program participation. Readers may be specifically concerned 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 which caused their disability. Thus, we cannot simply control for people's listed occupations and industries. However, we do construct several aggregate variables which can be used to alleviate the most obvious occupation-related concerns with our identification strategy.

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 rates. Specifically, we divide the number of injuries in the occupation by the number of private sector workers in the occupation. 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. Data on the number of workers in each occupation are obtained from the Occupational Employment Statistics. After assigning to each employed person in our sample injury rates for his or her occupation, we then construct average injury rates for each country of origin-PUMA cell. Similar measures are constructed for specific types of on-the-job injuries: sprains, chemical burns, and back pain. Descriptive statistics on these variables can be found in Table A1 of our Appendix.

Tables 4A and 4B present results from models which include controls for occupational hazards. Sample sizes are smaller in these specifications because there are some country of origin-PUMA 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). Baseline regressions run on this smaller sample yield almost identical results to regressions run on the full sample. As can be seen in Table 4A, immigrants residing in areas where people from their country of origin tend to work in jobs with high injury rates are more likely to receive DI. This is true when considering all injuries taken together as well as specific injuries. The prevalence of chemical burns in occupations typically held by local co-ethnics is highly correlated with DI participation, for example. However, adding controls for occupation-based injury rates to the DI model has virtually no effect on our measure of the importance of networks.

As discussed above, in order to qualify for SSI, applicants must be disabled and satisfy the income and asset tests. People with greater opportunities to work in more dangerous occupations may be more likely to become injured on the job thereby increasing their likelihood of satisfying the disability requirement. On the other hand, because riskier jobs tend to pay more (Leeth and Ruser 2003), they are less likely to satisfy the income and asset requirements. Which effect dominates is an empirical question, but our results, shown in Table 4B, suggest that the second effect is more important. In all specifications, there is a negative relationship between average on-the-job injury rates among co-ethnics living in a person's PUMA and the likelihood that that person is receiving SSI. However, the inclusion of these injury-related controls has no impact on the estimated coefficients on our network interactions.

Another potential issue related to the occupational distribution of immigrants is that immigrants from certain countries residing in specific areas may be more likely to have lost their jobs. To explore this labor market avenue, we construct country of origin-PUMA unemployment rates. Again, descriptive statistics are in Table A1 of our Appendix. As seen in the last columns of Tables 4A and 4B, our unemployment controls are positively associated with disability program participation, but our estimated network coefficients do not change when this variable is added to the model. Although results are not reported, we also ran regressions controlling for wages at various points in the wage distribution for each country of origin-PUMA cell. Regardless of whether we computed wages at the 10^{th}, 50^{th}, or 90^{th} percentile, higher wages were associated with lower disability program participation but the inclusion of these variables in the models had no impact on our estimated network coefficients.[17]

We conclude from this analysis that although occupational choice may be a strong predictor of disability program participation, the country of origin and PUMA fixed effects are already controlling for most of the variation in these variables.

[17] Wages at lower percentiles are especially important for DI recipients because the value of benefits relative to lost wages is substantially higher for low wage earners. As discussed in detail in Autor and Duggan (2003), this is because the DI benefit formula is progressive. Given that the benefit formula is indexed to the mean wage in the overall economy and low wage earners have experienced less wage growth than the average, their replacement rates have grown rather substantially over time. Moreover, low wage earners are less likely to receive health insurance from their employers making DI benefits, which automatically include Medicare after two years, even more attractive. Given that 85 percent of DI applicants were employed within the three years prior to applying while the comparable number for SSI applicants was only 30 percent (Bound, Burkhauser and Nichols 2001), wages are likely to be less influential for SSI applicants than DI applicants.

B. Networks and Ethnic Cohesion

For further evidence that our estimated network effects are actually measuring networks as opposed to omitted variables, we explore whether our estimated effects are indeed larger for people we would expect to be more socially connected to their groups. First, we separate the sample by English speaking ability. The first column in Tables 5A and 5B reports results from regressions run on a sample of immigrants who either do not speak English well or do not speak English at all while the second column reports results for a sample of immigrants who either can only speak English, speak English very well, or speak English well. In line with our expectations, immigrants fluent in English are less sensitive to ethnic networks when it comes to both DI and SSI participation. The language ability differential is even 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.

Next, we aim to compare network effects for the foreign born to the native born. From a theoretical perspective, the comparison is not clear-cut. On the one hand, we may expect the foreign born to have stronger ties to their ethnic group and so should be more sensitive to any ethnicity-based norms and taboos. They should also benefit more from information sharing about US programs than their native-born counterparts. On the other hand, the native born are significantly more likely than the foreign born to be eligible for disability programs. Thus, it is rather unclear whether network effects will be stronger for the foreign born or native born, but a finding that network effects are stronger for the foreign born could be viewed as strong evidence in favor of the importance of ethnic networks.

We cannot make the native-foreign comparison with the country of birth definition of ethnic origin. Instead, we define ethnicity by the first ancestry listed in the Census. Note that people who identify with a particular ancestry are very heterogeneous. They can range from second-generation immigrants whose parents arrived in the US shortly before their births to people whose families have been in the US for many generations. We also note that the native born who choose to write down an ancestry in the Census are likely to be more similar to people in their ethnic groups than the native born who do not identify with a particular ancestry (see Duncan and Trejo (2011) for a more formal exploration of this issue).

Tables 5A and 5B compare ancestry network effects for the native and foreign born. In both the DI and SSI models, results point to strong ancestry-based network effects for the foreign born but small effects for the native born which are statistically insignificant in the DI model. The fact that we do not see meaningful network effects for the native born suggests that natives either do not need the information provided by their ethnic networks or are not sensitive to any taboos within their ethnic networks.

C. Bureaucratic Channel

As discussed in Bertrand et al. (2000), bureaucracies can provide another potential explanation for our results which is unrelated to social interactions within ethnic networks. For example, local DDS offices may hire agents who speak a specific language whenever the number of people in the area that speak that language is sufficiently high.

Following Bertrand et al., we examine this possibility by restricting our sample to Spanish speakers. From the perspective of DDS managers, decisions about whether or not to hire Spanish speaking agents should depend on the total number of Spanish speakers without regard to country of origin. In contrast, if conditional on speaking the same language, Spanish-speaking immigrants are more likely to befriend immigrants from their country of birth, then we should be able to uncover network effects even in a sample restricted to immigrants from Spanish-speaking countries. [18]

As can be seen in the last column of Tables 5A and 5B, the estimated Spanish-only network coefficient in the DI model is about the same, both in magnitude and statistical significance, as that in the baseline specification, but the Spanish-only network coefficient in the SSI specification is not statistically different from zero. This may be either because network effects are simply not very important for SSI take-up among Spanish speakers or because there is not enough variation in average SSI receipt across Spanish-speaking countries to identify an effect.

7. How do Networks Operate?

Having provided evidence that social interactions play an important role in immigrants' disability program take-up, in this section we explore how. To use the terminology of Manski (1993), our analysis thus far has focused on distinguishing *correlated effects* from *exogenous/contextual* and *endogenous effects*. Correlated effects are a result of unobserved characteristics that affect individuals in a group simultaneously. These within group correlations would exist even if group members never came in contact with each other. 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.

Our empirical strategy does not allow us to distinguish between the different types of causal relationships. However, knowing the mechanisms through which

[18] These are Argentina, Bolivia, Chile, Colombia, Costa Rica, Cuba, Dominican Republic, Ecuador, El Salvador, Guatemala, Honduras, Mexico, Nicaragua, Panama, Paraguay, Peru, Spain, Uruguay and Venezuela.

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). Moreover, even among the types which do not generate multiplier effects, some can be used as evidence that disability benefits are not being awarded fairly while others are perfectly consistent with fairly awarded benefits. Although we are not able to perfectly distinguish between the mechanisms driving our network results, in this section we present several pieces of evidence which tend to be more consistent with some mechanisms than others.

A. Cultural Norms, Information Sharing, and Leisure Complementarities

The most often discussed sources of endogenous effects are cultural norms, information sharing, and leisure complementarities. All three would imply that a policy increasing the proportion of the population receiving benefits by one percentage point—by decreasing the threshold on the necessary severity of disability, for example—would ultimately increase the proportion of people receiving benefits by more than one percentage point. Evidence of these types of social interactions might also suggest that disability benefits are not being awarded fairly. If taboos and leisure complementarities are driving our results, then it is likely that some of the disability program applicants are in fact capable of working. On the other hand, if information sharing is important, then it may be that some of the people that are eligible for benefits do not receive them. We explore each of these mechanisms in turn, starting with social norms.

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. More importantly, as disability benefit take-up increases within a group, claiming benefits is likely to become even less stigmatized because of feedback effects. To examine whether social norms are likely to play an important role in disability program take-up, we replace our measure of network quality, average disability program usage in the ethnic group, with a more direct measure of attitudes toward unjustified take-up of social programs. Using various waves of the World Values Survey, we construct for each origin country the proportion of people who believe that "Claiming government benefits to which you are not entitled" is never justifiable. Values of this variable are shown for countries with the largest and smallest proportions in Appendix, Table A2. We merged this variable by country of origin to the immigrants in our sample and used it to replace the proportion of same country of origin immigrants that participate in the disability programs.[19] The results shown in Tables 6A and 6B suggest that an increase in the contact availability of co-ethnics has a smaller effect for people from countries where more people believe that claiming government benefits unfairly is never justifiable. This is certainly consistent with taboos and norms explaining our baseline findings.

Another attractive feature of using the World Values Survey variables is that it can alleviate some concerns regarding the Manski reflection problem. First, unlike the average disability take-up variable, the social norm variable is created from a very different sample than the one used in our analysis. The question regarding claiming government benefits is not even related to disabilities per se. The fact that we find evidence of network effects using this variable provides evidence that our analysis is not simply picking up within-group correlations in labor market opportunities or the tendency to become disabled.

Next, we examine whether information sharing within ethnic groups is likely to be a driving force behind our estimated network effects. A large literature

[19] We used the most recent wave of the World Values Survey from each country to construct this variable. Out of the 95 countries of birth in the original sample, we were able to construct the variable for 60 countries. To confirm that country selection was not driving our results, we ran the baseline models using only these 60 countries, and results did not change substantially.

documents the fact that many of the people eligible for social assistance programs do not apply for them (see Aizer 2007 for references). Aizer (2007) finds that access to bilingual application assistants increases Medicaid enrollment rates of Hispanic and Asian children suggesting that lack of information about program eligibility and how to enroll may explain low take-up rates. There is also evidence that ethnic networks aided in the transmission of new information about Special Supplemental Nutrition Program for Women, Infants and Children (WIC) eligibility among pregnant Hispanic women during the period surrounding welfare reform (Figlio, Hamersma, and Roth 2011). In contrast, however, information sharing does not seem to be a major driver of the estimated network effects in the take-up of publically funded prenatal care (Aizer and Currie 2004) or welfare dependence (Aslund and Fredriksson 2009). [20]

Although we do not have a natural experiment which might be used to clearly differentiate the information story from other potential drivers of network effects, we can examine whether the people who are likely to be at the greatest information disadvantages increase disability take-up most when they reside amidst others who are receiving disability benefits. More specifically, we split the sample based on educational attainment and examine whether estimated network effects are largest for immigrants with the least amount of education.

As can be seen in the DI specifications of Table 6A, splitting the sample by educational attainment does not generate a clear pattern regarding network effects for immigrants with less than a college degree. The estimated network coefficient is rather large but statistically insignificant for immigrants without a high school

[20] Aizer and Currie (2004) conclude that information sharing is not likely to be important because estimated ethnic network effects are large even among women who have already used the program for prior births. Aslund and Fredriksson (2009) make a similar conclusion in the context of welfare dependence among refugees in Sweden because all of the refugees in their sample are introduced to welfare upon arrival. Figlio et al. (2011) suggest that while information sharing may not be important in the context of rather stable programs, it may be important when there is a real or perceived change in a program's eligibility requirements or application process as was the case during welfare reform.

degree, smaller in magnitude but statistically significant for high school graduates, and largest and statistically significant for those who have completed some years of college. The estimated network effects are considerably smaller in magnitude and not statistically significant for college-educated immigrants. This last result may be interpreted as evidence of information sharing given that college graduates should be well-equipped to navigate the entire DI application process without requiring help from social contacts. It is also possible that norms are the only mechanism through which networks matter but that the college-educated are simply not as sensitive to the norms within their ethnic groups.

The estimated network effects in SSI specifications, shown in Table 6B, display a much clearer pattern in that estimated interaction coefficients are larger for immigrants with fewer years of schooling. These results are certainly more consistent with an information story than the DI results just as one might expect given that information provided through networks may be especially important for the population qualifying for SSI benefits.

As an additional test of the information sharing story, we add to our baseline DI specification an interaction between contact availability and the proportion of co-ethnics receiving SSI and to our baseline SSI specification an interaction between contact availability and the proportion of co-ethnics receiving DI. The DI and SSI programs use the same rules for determining who is disabled and have the same system of appeals. This implies that if information sharing is the driving force behind our estimated network effects, we might expect that conditional on average DI take-up in a person's ethnic group, immigrants in groups with more SSI take-up would be more likely to receive DI benefits when they reside amidst many others from their own ethnic group. Similarly, conditional on average co-ethnic SSI take-up, immigrants in groups with higher DI take-up should be more likely to receive SSI benefits when they reside amidst co-ethnics. It turns out, however, that results, shown in column 6 of Tables 6A and 6B, are exactly opposite from this prediction. This certainly cannot be taken as definitive proof against the information sharing hypothesis, but it may suggest that there are other mechanisms at play.

Finally, we 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. In fact, a person with a marginal disability might even enjoy leisure more with a person who is healthy but unable or unwilling to find employment than with a person who is incapable of working. Thus, we add to our baseline models an interaction between contact availability and the percentage of co-ethnics that are not employed. Values of this variable are shown for origin countries with the largest and smallest proportions not employed in the Appendix, Table A2.

As can be seen in the last column of Tables 6A and 6B, 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. 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 program (Borghans, Gielen, and Luttmer 2010; Lindner 2011). 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 not employed interactions are added to the models.

We conclude from these informal tests that while cultural norms may play an important role in explaining our network results, leisure complementarities are not likely to be driving results. Information sharing may play some role in explaining network effects in SSI take-up, but it does not appear that information sharing is a major driver of network effects in DI take-up.

B. Other Social Interaction Effects

Social interactions may also have positive causal impacts on behaviors for reasons not directly related to disability program participation. In fact, social interactions need not imply a multiplier effect and may be perfectly consistent with disability benefits being awarded fairly. For example, friends and family members may impact ultimate disability program participation through their influence on people's health behaviors. If there are peer effects in determining smoking rates, as suggested by Fletcher (2010), then immigrants belonging to ethnic groups with high smoking rates may be more likely to smoke if they reside amidst many others from their ethnic group. To the extent that smoking has a causal impact on ultimate disability and hence the take-up of disability programs, then our network results could be explained completely by social interactions in the transmission of disability.[21] The US Census does not contain information on health behaviors and so we are not able to test this hypothesis directly. Census-responders were asked whether they had a physical or mental condition which "causes difficulty working, limits the amount or type of work they can do, or prevents them from working altogether." As shown in the first column of Tables 7A and 7B, adding this variable to our baseline specification slightly decreases the magnitude of our estimated network coefficients but they remain statistically and economically

[21] An analogous story could be told if there are peer effects in obesity as suggested by Fowler and Christakis (2008). It should be noted, however, that Cohen-Cole and Fletcher (2008) fail to find peer effects in obesity determination when standard econometric techniques are used to control for exogenous effects. Similarly, there are strong within friendship network correlations in depression (Rosenquist, Fowler and Christakis 2011), but when exploiting exogenous variation arising from college roommate assignment, Eisenberg, Golberstein, Whitlock and Downs (2011) find only modest peer effects for depression, and even those small effects are found only for men.

significant suggesting that health-related behaviors are not likely to be driving our results.

Next we consider the role of networks in the determination of labor market opportunities of immigrants. There is a large literature documenting how personal connections aid in finding jobs (Munshi 2003; Bayer, Ross and Topa 2008; Cappellari and Tatsiramos 2011). If immigrants in ethnic groups with more labor market success are better able to find lucrative jobs when they reside amidst others from the same ethnic group, then people with marginal disabilities may find it optimal to continue working despite hardship. A parallel literature presents evidence of networks in welfare take-up (Bertrand et al. 2000; 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 constraints for SSI, our estimated network effects may simply reflect the role of social interactions in determining who qualifies for the disability programs.

To examine this possibility, we exploit the fact that the disability and retirement programs of the Social Security Administration have almost the same eligibility requirements. To qualify for Social Security retirement income, individuals must satisfy the same work history requirements as DI-recipients (they need not satisfy the recent work requirement) but receive benefits, irrespective of disability, as long as they are above a certain age. Given the magnitude of the program,[22] 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

[22] In 2008, Social Security paid benefits to 88 percent of married couples and 86 percent of unmarried individuals aged 65 and above (Annual Statistical Supplement to the Social Security Bulletin 2010).

straightforward. Thus, a positive and statistically significant interaction coefficient in a model with the receipt of Social Security retirement income as the dependent variable might be interpreted as evidence of the role of networks in terms of satisfying work history requirements as opposed to applying for or receiving benefits.

Similarly, SSI is available to individuals age 65 and above, regardless of disability status, as long as applicants meet the income and asset requirements. Findings that networks help determine SSI-receipt for this population would suggest that social contacts create or maintain a culture of poverty which makes people eligible for SSI, for reasons unrelated to disability. We note, however, that the jump in the percentage of SSI recipients after age 65, from 1.3 to 10.4 in our sample of immigrants, is significantly lower than the comparable jump for DI to retirement income of 1.6 to 69.9. Thus, information sharing and taboos may be important determinants of receiving SSI even for people above retirement age.

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. To measure Social Security retirement income, we run the baseline specifications (equation 1) on individuals age 65 and above. Again, for people below retirement age, they must have a disability in order to receive benefits, but anyone who meets the income and asset constraints can receive SSI if they are above retirement age.

Tables 7A and 7B (Columns 2-7) show results of our models run on an age 65 plus sample of non-workers. 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 about a third the size of the network coefficient in the DI model while the retirement-age SSI network coefficient is 43 percent smaller than the comparable coefficient in the baseline SSI model. This suggests that similarities in eligibility can explain part, but not all, of the estimated network effects in our baseline models.

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 Table A3 of our Appendix, the estimated network coefficient for DI in the 57-61 year old Census sample is 0.104 with a p-value of 0.372. In the retirement-age ACS sample, the estimated DI network coefficient is 0.015 with a p-value of 0.613. Neither coefficient is statistically significant, potentially because of the relatively small sample sizes, but it is quite telling that the magnitude drops so substantially just after retirement age. In the SSI models (Table A4), the network coefficient dropped from 0.616 with a p-value of 0.000 to 0.227 with a p-value of 0.000. 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 and its square in all of our 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 actually positive, although not statistically significant, suggesting that if anything, our retirement sample results should overestimate eligibility effects. Results are shown in Tables A3, A4 of the Appendix.

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 samples. While norms and taboos are likely to play a large role in determining who exaggerates disabilities, they are unlikely to be a big factor in determining who receives Social Security retirement benefits or SSI for people age 65 and above.

Using the retirement age sample, we test whether residing amidst many co-ethnics makes people especially more likely to receive Social Security retirement income (or SSI) if they are from countries where fewer people believe it is always unethical to fraudulently receive government benefits. Results, shown in column 3 of Tables 7A and 7B suggest that social norms have no statistical or economic impact on the receipt of benefits in our older samples. This is in direct contrast to results in the baseline samples, making us more confident that our empirical strategy used on the working age sample is at least partially identifying actual network effects which are very likely to operate via social norms.

Next, we explore whether the relationships we found in the working age samples between education and estimated network effects disappear when using the retirement age sample. Because the older immigrants need not prove the existence of a disability, information sharing should be significantly less important in this sample, and so differences in network effects by education level should be much less pronounced.

Recall that Table 6A (columns 2-5) showed no evidence of network effects for DI among working-age immigrants with a college degree, but among those with less than a college degree, the estimated network effect did not vary very much with educational attainment. As can be seen in Table 7A (columns 4-7), the same pattern emerges in the retirement age sample suggesting that differences in eligibility for DI can explain the education pattern found in our working age sample. This may be because, among immigrants with less than a college degree, workers with more experience are more likely to live around co-ethnics if co-ethnics have more years of experience while, among college graduates, this relationship does not hold.

A different story can be told with respect to the SSI models shown in Table 7B. While in the working age sample, there is a clear negative relationship between educational attainment and estimated network effects, estimated network effects do not vary with education in the retirement age sample. This suggests that income and asset restrictions cannot explain the education patterns seen in the working age sample. From this analysis, we conclude that while norms may play an important role in explaining both DI and SSI take-up, information sharing is unlikely to be an important determinant of DI take-up but may be important for SSI take-up.

8. Conclusion

Although we do not claim to perfectly identify the role of networks in any one specification, we believe that taken together, our analyses make a strong case for the conclusion that networks play a large role in determining who receives disability payment. The stated aim of the Social Security Disability Insurance program is to insure below retirement age workers against the risk of not being able to perform “substantial” work due to a physical or mental disability. Our finding that networks play a large role in determining who receives disability payments suggests that the Social Security Administration is not doing a fair and effective job of allocating disability insurance funds. Census data do not allow us to formally decompose the mechanisms driving our results, but the evidence we provide is consistent with ethnic network effects being driven by social norms, inconsistent with leisure complementarities, and partially consistent with information sharing. We also show 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.

We view our results as suggestive of how social interactions affect disability program take-up in general but our analysis focuses on immigrants. Information sharing within networks is likely to be more important for the foreign born than for natives. 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	4.47	611	Cambodia	7.34	3,940
Azores	4.02	808	Laos	6.18	5,576
Croatia	3.87	816	Belarus	4.37	511
Italy	3.53	9,480	Armenia	4.19	1,321
Portugal	3.48	5,808	Ukraine	3.77	3,342
Bottom 5			Bottom 5		
Taiwan	0.44	8,483	South Africa	0.22	1,213
Sri Lanka	0.42	565	Indonesia	0.19	1,200
Burma	0.41	783	Liberia	0.18	632
Belgium	0.29	535	Switzerland	0.11	719
New Zealand	0.25	505	Norway	0.00	554

Notes: Our sample consists of non-widowed, non-institutionalized immigrants, age 25 to 61, who were living in the US five years prior to the survey. Only countries with more than 500 observations are considered.

Table 2 Summary Statistics

	Whole Sample		DI Sample	SSI Sample
	(1)	(2)	(3)	(4)
	Mean	Std. Dev.	Mean	Mean
DI	0.016	0.126	---	0.155
SSI	0.013	0.112	0.122	---
Age	40.969	9.64	46.962	46.295
Male	0.511	0.499	0.522	0.46
High school dropout	0.326	0.469	0.457	0.550
High school degree	0.301	0.459	0.309	0.292
Some college	0.145	0.352	0.117	0.091
English fluency	0.481	0.499	0.401	0.322
Spouse present	0.686	0.464	0.602	0.479
Child	0.643	0.479	0.583	0.537
Number of children	2.217	1.191	2.157	2.363
Hispanic	0.229	0.421	0.238	0.217
Black	0.071	0.257	0.086	0.086
Asian	0.248	0.431	0.168	0.253
Other race	0.004	0.063	0.003	0.003
Years in the US	18.598	10.352	22.585	20.15
Years in the US^2/100	4.53	4.848	6.433	5.175
Contact availability (CA) in levels	0.069	0.103	0.081	0.084
CA	-4.093	2.015	-3.866	-3.717
Observations	692066		11509	8922

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. 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). 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 "Number of children" refers to the total number of children in the household. "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".

Table 3A Effects of Networks on Disability Insurance Receipt

Dependent variable: DI			
	(1)	(2)	(3)
CA * Proportion of co-ethnics receiving DI	0.124**	0.107**	0.103**
	(0.023)	(0.023)	(0.023)
CA	-0.001**	-0.001*	-0.001**
	(0.000)	(0.000)	(0.000)
Male		0.001*	0.001**
		(0.000)	(0.000)
Married, Spouse present		-0.008**	-0.008**
		(0.000)	(0.000)
Child		0.001	0.001+
		(0.001)	(0.001)
Number of children		-0.0001	-0.0005*
		(0.000)	(0.000)
Hispanic		0.001	0.001
		(0.001)	(0.001)
Black		0.007**	0.007**
		(0.002)	(0.002)
Asian		-0.001	-0.0002
		(0.001)	(0.001)
Other race		-0.002	-0.003
		(0.002)	(0.002)
High school dropout			0.013**
			(0.001)
High school degree			0.007**
			(0.000)
Some college			0.004**
			(0.001)
English fluency			-0.00**
			(0.000)
Years in the US			0.0002*
			(0.000)
Years in the US^2/100			0.0002
			(0.000)
Country fixed effects	Yes	Yes	Yes
PUMA fixed effects	Yes	Yes	Yes
Age fixed effects	No	Yes	Yes
Multiplier	0.324	0.293	0.287
Observations	692,066	692,066	692,066
R-squared	0.008	0.017	0.018

Notes: See Table 1 notes for information on the sample and Table 2 for notes on the variables. The omitted education dummy is "College and more". The omitted race dummy is "white". Heteroskedasticity corrected standard errors clustered by country of origin and PUMA (66364 cells) are in parentheses. Observations are weighted using the appropriate person-level weights provided by the 2000 U.S. Census. Significance levels are noted by the following: ** significance at 1%, *significance at 5%,
+ significance at 10%. A description of how to calculate the "multiplier" is provided in the text.

Table 3B Effects of Networks on Supplemental Security Income Receipt

Dependent variable: SSI			
	(1)	(2)	(3)
CA * Proportion of co-ethnic receiving SSI	0.285**	0.275**	0.279**
	(0.031)	(0.030)	(0.030)
CA	-0.003**	-0.002**	-0.003**
	(0.000)	(0.000)	(0.000)
Male		-0.002**	-0.002**
		(0.000)	(0.000)
Married, Spouse present		-0.013**	-0.013**
		(0.000)	(0.000)
Child		-0.002**	-0.001*
		(0.001)	(0.001)
Number of children		0.001**	0.0002
		(0.000)	(0.000)
Hispanic		-0.001+	-0.002**
		(0.000)	(0.000)
Black		0.003*	0.004*
		(0.002)	(0.002)
Asian		-0.001	-0.001
		(0.001)	(0.001)
Other race		-0.001	-0.002
		(0.003)	(0.003)
High school dropout			0.019**
			(0.001)
High school degree			0.009**
			(0.000)
Some college			0.004**
			(0.000)
English fluency			-0.002**
			(0.000)
Years in the US			0.0003**
			(0.000)
Years in the US^2/100			-0.001**
			(0.000)
Country fixed effects	Yes	Yes	Yes
PUMA fixed effects	Yes	Yes	Yes
Age fixed effects	No	Yes	Yes
Multiplier	0.517	0.508	0.513
Observations	692,066	692,066	692,066
R-squared	0.016	0.024	0.027

Notes: See Table 1 notes for information on the sample and Table 2 for notes on the variables. The omitted education dummy is “College and more”. The omitted race dummy is “white”. Heteroskedasticity corrected standard errors clustered by country of origin and PUMA (66364 cells) are in parentheses. Observations are weighted using the appropriate person-level weights provided by the 2000 U.S. Census. Significance levels are noted by the following: ** significance at 1%, * significance at 5%, + significance at 10%. A description of how to calculate the “multiplier” is provided in the text.

Table 4A Effect of Occupational Injuries and Unemployment on DI

Dependent variable: DI	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
CA * Proportion of co-ethnics receiving DI	0.102**	0.102**	0.102**	0.102**	0.102**	0.102**	0.102**	0.102**
	(0.023)	(0.023)	(0.023)	(0.023)	(0.023)	(0.023)	(0.023)	(0.023)
CA	-0.001**	-0.001**	-0.001**	-0.001**	-0.001**	-0.001**	-0.001**	-0.001**
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)
On-the-job injuries in country of origin-PUMA cells	0.116*							
	(0.057)							
On-the-job fractures in country of origin-PUMA cells		1.138+						
		(0.656)						
On-the-job bruises in country of origin-PUMA cells			0.999+					
			(0.560)					
On-the-job sprains in country of origin-PUMA cells				0.308*				
				(0.133)				
On-the-job chemical burns in country of origin-PUMA cells					21.608**			
					(6.869)			
On-the-job pain in country of origin-PUMA cells						1.432*		
						(0.665)		
On-the-job back pain in country of origin PUMA-cells							4.001*	
							(1.725)	
Unemployment rate in country of origin-PUMA cells								0.006*
								(0.003)
Country 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
Observations	684,979	684,979	684,979	684,979	684,979	684,979	684,979	692,066
R-squared	0.018	0.018	0.018	0.018	0.018	0.018	0.018	0.018

Notes: All control variables shown in Table 3A are included in these models. See Table 1 notes for information on the sample and Table 2 for notes on the variables. Heteroskedasticity corrected standard errors clustered by country of origin and PUMA are in parentheses. Observations are weighted using the appropriate person-level weights provided by the 2000 U.S. Census. The number of observations in Columns 1 to 7 is less than in Column 8 because we were not able to merge in data from the Bureau of Labor Statistics' (BLS) injuries, illnesses and fatalities program for everyone in the baseline sample. Significance levels are noted by the following: ****** significance at 1%, * significance at 5%, + significance at 10%.

Table 4B Effect of Occupational Injuries and Unemployment on SSI

Dependent variable: SSI	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
CA * Proportion of co-ethnics receiving SSI	0.283**	0.284**	0.283**	0.284**	0.284**	0.284**	0.284**	0.280**
	(0.031)	(0.031)	(0.031)	(0.031)	(0.031)	(0.031)	(0.031)	(0.030)
CA	-0.003**	-0.003**	-0.003**	-0.003**	-0.003**	-0.003**	-0.003**	-0.003**
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)
On-the-job injuries in country of origin-PUMA cells	-0.118*							
	(0.053)							
On-the-job fractures in country of origin-PUMA cells		-1.054+						
		(0.599)						
On-the-job bruises in country of origin-PUMA cells			-1.246*					
			(0.529)					
On-the-job sprains in country of origin-PUMA cells				-0.239*				
				(0.116)				
On-the-job chemical burns in country of origin-PUMA cells					-6.573			
					(7.592)			
On-the-job pain in country of origin-PUMA cells						-1.048+		
						(0.587)		
On-the-job back pain in country of origin PUMA-cells							-1.913	
							(1.493)	
Unemployment rate in country of origin-PUMA cells								0.007**
								(0.002)
Country 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
Observations	684,979	684,979	684,979	684,979	684,979	684,979	684,979	692,066
R-squared	0.027	0.027	0.027	0.027	0.027	0.027	0.027	0.027

Notes: All control variables shown in Table 3A are included in these models. See Table 1 notes for information on the sample and Table 2 for notes on the variables. Heteroskedasticity corrected standard errors clustered by country of origin and PUMA are in parentheses. Observations are weighted using the appropriate person-level weights provided by the 2000 U.S. Census. The number of observations in Columns 1 to 7 is less than in Column 8 because we were not able to merge in data from the Bureau of Labor Statistics' (BLS) injuries, illnesses and fatalities program for everyone in the baseline sample. Significance levels are noted by the following: ** significance at 1%, * significance at 5%, + significance at 10%.

Table 5A Robustness Checks, DI Model

	English-speaking ability		Ancestry-defined measure of ethnicity		Spanish speakers
	Not fluent	Fluent	Foreign born	Native born	
Dependent variable: DI	(1)	(2)	(3)	(4)	(5)
CA* Proportion of co-ethnics receiving DI	0.093*	0.061*	0.082**	-0.024	0.087*
	(0.040)	(0.024)	(0.024)	(0.015)	(0.039)
CA	-0.001	-0.0004	-0.001**	0.00004	-0.001+
	(0.001)	(0.000)	(0.000)	(0.000)	(0.001)
Country fixed effects	Yes	Yes	Yes	Yes	Yes
PUMA fixed effects	Yes	Yes	Yes	Yes	Yes
Age fixed effects	Yes	Yes	Yes	Yes	Yes
Observations	361,524	330,542	679,472	3,823,401	347,268
R-squared	0.023	0.022	0.021	0.028	0.022

Notes: All control variables shown in Table 3A are included in these models. See Table 1 notes for information on the sample and Table 2 for notes on the variables. Heteroskedasticity corrected standard errors are in parentheses. They are clustered by country of origin and PUMA in columns 1, 2, and 5 but on ancestry and PUMA in columns 3 and 4. Observations are weighted using the appropriate person-level weights provided by the 2000 U.S. Census. Significance levels are noted by the following:
** significance at 1%, * significance at 5%, + significance at 10%.

Table 5B Robustness Checks, SSI Model

	English-speaking ability		Ancestry-defined measure of ethnicity		Spanish speakers
	Not fluent	Fluent	Foreign born	Native born	
Dependent variable: SSI	(1)	(2)	(3)	(4)	(5)
CA* Proportion of co-ethnics receiving SSI	0.372**	0.056**	0.271**	0.028*	0.034
	(0.040)	(0.021)	(0.025)	(0.011)	(0.026)
CA	-0.005**	-0.0002	-0.004**	-0.001**	-0.0004
	(0.001)	(0.000)	(0.000)	(0.000)	(0.000)
Country fixed effects	Yes	Yes	Yes	Yes	Yes
PUMA fixed effects	Yes	Yes	Yes	Yes	Yes
Age fixed effects	Yes	Yes	Yes	Yes	Yes
Observations	361,524	330,542	679,472	3,823,401	347,268
R-squared	0.038	0.020	0.038	0.036	0.023

Notes: All control variables shown in Table 3A are included in these models. See Table 1 notes for information on the sample and Table 2 for notes on the variables. Heteroskedasticity corrected standard errors are in parentheses. They are clustered by country of origin and PUMA in columns 1, 2, and 5 but on ancestry and PUMA in columns 3 and 4. Observations are weighted using the appropriate person-level weights provided by the 2000 U.S. Census. Significance levels are noted by the following:
** significance at 1%, * significance at 5%, + significance at 10%.

Table 6A Mechanisms Through Which Networks Operate: DI

	Social Norms	Information Sharing					Leisure Complementarities
		HS Dropout	HS Degree	Some College	College		
Dependent variable: DI	(1)	(2)	(3)	(4)	(5)	(6)	(7)
CA * Proportion of co-ethnics receiving DI		0.082	0.080*	0.087*	0.041	0.129**	0.131**
		(0.067)	(0.037)	(0.043)	(0.033)	(0.025)	(0.024)
CA * Origin country beliefs	-0.037**						
	(0.013)						
CA * Proportion of co-ethnics receiving SSI						-0.0361*	
						(0.014)	
CA * Proportion of co-ethnics not employed							-0.009**
							(0.002)
CA	0.002**	-0.001	-0.001	-0.001	-0.0001	-0.001*	0.001*
	(0.000)	(0.001)	(0.001)	(0.001)	(0.001)	(0.000)	(0.001)
Country fixed effects	Yes	Yes	Yes	Yes	Yes	Yes	Yes
PUMA fixed effects	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Age fixed effects	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Observations	536,407	231,850	207,132	98,923	154,161	692,066	692,066
R-squared	0.019	0.034	0.024	0.035	0.027	0.018	0.018

Notes: "Origin country beliefs" refers to the proportion of people in a person's home country who believe that "Claiming government benefits to which you are not entitled" is never justifiable. All control variables shown in Table 3A are included in these models. See Table 1 notes for information on the sample and Table 2 for notes on the control variables. Heteroskedasticity corrected standard errors clustered by country of origin and PUMA are in parentheses. Observations are weighted using the appropriate person level weights provided by the 2000 U.S. Census. Significance levels are noted by the following: ** significance at 1%, * significance at 5%, + significance at 10%.

Table 6B Mechanisms Through Which Networks Operate: SSI

	Social Norms	Information Sharing					Leisure Complementarities
		HS Dropout	HS Degree	Some College	College		
Dependent variable: SSI	(1)	(2)	(3)	(4)	(5)	(6)	(7)
CA * Proportion of co-ethnics receiving SSI		0.512**	0.146**	0.056*	0.073**	0.335**	0.318**
		(0.058)	(0.029)	(0.023)	(0.019)	(0.033)	(0.031)
CA * Origin country beliefs	-0.036**						
	(0.014)						
CA * Proportion of co-ethnics receiving DI						-0.207**	
						(0.027)	
CA * Proportion of co-ethnics not employed							-0.021**
							(0.002)
CA	0.001**	-0.008**	-0.002**	0.0002	-0.001*	-0.001	0.002**
	(0.000)	(0.001)	(0.000)	(0.000)	(0.000)	(0.000)	(0.001)
Country fixed effects	Yes	Yes	Yes	Yes	Yes	Yes	Yes
PUMA fixed effects	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Age fixed effects	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Observations	536,407	231,850	207,132	98,923	154,161	692,066	692,066
R-squared	0.019	0.051	0.025	0.036	0.022	0.027	0.027

Notes: “Origin country beliefs” refers to the proportion of people in a person’s home country who believe that “Claiming government benefits to which you are not entitled” is never justifiable. All control variables shown in Table 3A are included in these models. See Table 1 notes for information on the sample and Table 2 for notes on the control variables. Heteroskedasticity corrected standard errors clustered by country of origin and PUMA are in parentheses. Observations are weighted using the appropriate person-level weights provided by the 2000 U.S. Census. Significance levels are noted by the following: ** significance at 1%, * significance at 5%, + significance at 10%.

Table 7A Other Social Interaction Effects: DI

	Baseline Sample	Retirement Age Sample					
		Main specification	Cultural Norms	Information Sharing			
				HS Dropout	HS Degree	Some College	College
Dependent variable: DI	(1)	(2)	(3)	(4)	(5)	(6)	(7)
CA * Proportion of co-ethnics receiving DI	0.104**	0.036**		0.020	0.021	0.081**	0.029
	(0.022)	(0.009)		(0.016)	(0.016)	(0.028)	(0.023)
CA * Origin country beliefs			0.001				
			(0.001)				
CA	-0.002**	-0.031**	-0.009**	-0.018+	-0.024*	-0.061**	-0.026+
	(0.000)	(0.006)	(0.002)	(0.011)	(0.011)	(0.020)	(0.015)
Disability dummy	0.013**						
	(0.001)						
Country fixed effects	Yes	Yes	Yes	Yes	Yes	Yes	Yes
PUMA fixed effects	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Age fixed effects	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Observations	692,066	86,963	71,456	35,885	27,133	9,491	14,454
R-squared	0.019	0.246	0.262	0.276	0.271	0.347	0.342

Notes: Column 1 uses our baseline sample of 25 to 61 year olds but controls for whether the respondent has "a physical or mental health condition that causes difficulty working, limits the amount or type of work they can do, or prevents them from working altogether". Columns 2 through 7 use a sample of people at or above the age of 65. "Origin country beliefs" refer to the proportion of people in a person's home country who believe that "Claiming government benefits to which you are not entitled" is never justifiable. All control variables shown in Table 3A are included in all the models in this table. See Table 1 notes for information on the sample and Table 2 for notes on the control variables. Heteroskedasticity corrected standard errors clustered by country of origin and PUMA are in parentheses. Observations are weighted using the appropriate person-level weights provided by the 2000 U.S. Census. Significance levels are noted by the following: ** significance at 1%, * significance at 5%, + significance at 10%.

Table 7B Other Social Interaction Effects: SSI

	Baseline Sample	Retirement Age Sample					
		Main specification	Cultural Norms	Information Sharing			
				HS Dropout	HS Degree	Some College	College
Dependent variable: SSI	(1)	(2)	(3)	(4)	(5)	(6)	(7)
CA * Proportion of co-ethnics receiving SSI	0.281**	0.160**		0.145**	0.181**	0.141**	0.129**
	(0.030)	(0.014)		(0.026)	(0.020)	(0.038)	(0.031)
CA * Origin country beliefs			-0.001				
			(0.001)				
CA	-0.003**	-0.015**	0.012**	-0.019**	-0.012**	-0.012*	-0.015**
	(0.000)	(0.002)	(0.002)	(0.004)	(0.003)	(0.005)	(0.004)
Disability dummy	0.019**						
	(0.001)						
Country fixed effects	Yes	Yes	Yes	Yes	Yes	Yes	Yes
PUMA fixed effects	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Age fixed effects	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Observations	692,066	86,963	71,456	35,885	27,133	9,491	14,454
R-squared	0.031	0.150	0.167	0.150	0.182	0.320	0.306

Notes: Column 1 uses our baseline sample of 25 to 61 year olds but controls for whether the respondent has "a physical or mental health condition that causes difficulty working, limits the amount or type of work they can do, or prevents them from working altogether". Columns 2 through 7 use a sample of people at or above the age of 65. "Origin country beliefs" refer to the proportion of people in a person's home country who believe that "Claiming government benefits to which you are not entitled" is never justifiable. All control variables shown in Table 3A are included in all the models in this table. See Table 1 notes for information on the sample and Table 2 for notes on the control variables. Heteroskedasticity corrected standard errors clustered by country of origin and PUMA are in parentheses. Observations are weighted using the appropriate person-level weights provided by the 2000 U.S. Census. Significance levels are noted by the following: ** significance at 1%, * significance at 5%, + significance at 10%.

Appendix

Table A1-Descriptive Statistics for On-the-Job Injury and Unemployment Rates within Country of Origin-PUMA cells

Variable	Mean	Std. Deviation
Total injuries	218.14	251.55
Fractures	10.02	5.33
Bruises	11.61	6.15
Sprains	52.56	22.90
Chemical burns	0.75	0.53
Pain	10.29	4.51
Back pain	3.56	1.72
Unemployment rate	0.06	0.08

Notes: Occupational injury rates, per 10,000 workers, were constructed using data on injuries by occupation from the Bureau of Labor Statistics (BLS) injuries, illnesses, and fatalities (IIF) program and data on number of workers by occupation from the Occupational Employment Statistics. These variables were merged with 2000 Census data by occupation and then averaged within country of origin-PUMA cells. On average, there are 218 on-the-job injuries within the occupations held by co-ethnics living within the PUMAs of individuals in our sample.

Table A2-Average Origin Country Beliefs about Claiming Government Benefits and Proportion of Co-ethnics Not Employed

	Origin Country Beliefs		Not Employed
Top 5		Top 5	
Bangladesh	0.92	Dominican Republic	0.44
Turkey	0.89	Mexico	0.42
Morocco	0.86	Armenia	0.40
Denmark	0.83	Cambodia	0.39
Jordan	0.79	Guatemala	0.39
Bottom 5		Bottom 5	
Belarus	0.35	Former Soviet Union, not specified*	0.19
Philippines	0.27	Liberia	0.19
Greece	0.24	Hong Kong	0.19
Malaysia	0.18	New Zealand	0.18
Thailand	0.14	Nigeria	0.16

Notes: Origin country beliefs refer to the proportion of people within a person's home country who believe that "Claiming government benefits to which you are not entitled" is never justifiable. This variable was constructed using data from the World Values Survey. The "not employed" variable was constructed using our Census sample. *The specified countries are Estonia, Latvia, Lithuania, Baltic States, Russia, Byelorussia, Moldavia, Bessarabia, Ukraine, Armenia, Azerbaijan, Georgia, Kazakhstan, Kirghizia, Tadzhik, Turkmenistan, Uzbekistan, and Siberia.

Table A3-Robustness Checks, DI Model

Dependent Variable: DI	Baseline Sample	Age 57-61 in 2000 Cohort	
		Census 2000	ACS 2008-2010
	DI	DI	Social Security
	(1)	(2)	(3)
CA * Proportion of co-ethnics receiving DI	0.042	0.104	0.015
	(0.045)	(0.117)	(0.031)
CA * Proportion of co-ethnics receiving DI * Years in the US	0.001		
	(0.002)		
Years in the US * CA	0.0001**		
	(0.000)		
Proportion of co-ethnics receiving DI * Years in the US	0.016		
	(0.012)		
CA	-0.002**	0.002	-0.004
	(0.001)	(0.002)	(0.022)
Country fixed effects	Yes	Yes	Yes
PUMA fixed effects	Yes	Yes	Yes
Age fixed effects	Yes	Yes	Yes
Observations	692,066	53,704	34,815
R-squared	0.018	0.066	0.276

Notes: Column 1 uses the baseline sample, 25-61 year olds, from the 2000 Census. Columns 2 and 3 use data on a cohort aged 57-61 in the year 2000. The dependent variable in column 2 is DI. The sample, from the 2000 Census, includes 57-61 year olds. The dependent variable in column 3 is Social Security retirement income. The data come from the ACS 2008-2010 samples of the same cohort; 65-69 year olds in 2008, 66-70 year olds in 2009, and 67-71 year olds in 2010. All control variables shown in Table 3A are included in all the models in this table. See Table 1 notes for information on the sample and Table 2 for notes on the control variables. Heteroskedasticity corrected standard errors clustered by country of origin and PUMA are in parentheses. Observations are weighted using the appropriate person-level weights provided by the 2000 U.S. Census. Significance levels are noted by the following: ** significance at 1%, * significance at 5%, + significance at 10%.

Table A4-Robustness Checks, SSI Model

Dependent Variable: SSI	Baseline Sample	Age 57-61 in 2000 Cohort	
		Census 2000	ACS 2008-2010
		SSI	Social Security
	(1)	(2)	(3)
CA * Proportion of co-ethnics receiving SSI	0.244**	0.616**	0.227**
	(0.033)	(0.111)	(0.032)
CA * Proportion of co-ethnics receiving SSI * Years in the US	0.002**		
	(0.001)		
Years in the US * CA	0.0001**		
	(0.000)		
Proportion of co-ethnics receiving SSI * Years in the US	0.012**		
	(0.003)		
CA	-0.004**	-0.007**	-0.009**
	(0.000)	(0.002)	(0.002)
Country fixed effects	Yes	Yes	Yes
PUMA fixed effects	Yes	Yes	Yes
Age fixed effects	Yes	Yes	Yes
Observations	692,066	53,704	34,815
R-squared	0.027	0.096	0.144

Notes: Column 1 uses the baseline sample, 25-61 year olds, from the 2000 Census. Columns 2 and 3 use data on a cohort aged 57-61 in the year 2000. The dependent variable in column 2 is DI. The sample, from the 2000 Census, includes 57-61 year olds. The dependent variable in column 3 is Social Security retirement income. The data come from the ACS 2008-2010 samples of the same cohort; 65-69 year olds in 2008, 66-70 year olds in 2009, and 67-71 year olds in 2010. All control variables shown in Table 3A are included in all the models in this table. See Table 1 notes for information on the sample and Table 2 for notes on the control variables. Heteroskedasticity corrected standard errors clustered by country of origin and PUMA are in parentheses. Observations are weighted using the appropriate person-level weights provided by the 2000 U.S. Census. Significance levels are noted by the following: ** significance at 1%, * significance at 5%, + significance at 10%.