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George J. Borjas
Glenn T. Sueyoshi

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ABSTRACT

There exist sizeable differences in the incidence and duration of welfare spells across ethnic groups, and these differences tend to persist across generations. Using the National Longitudinal Surveys of Youth, we find that children raised in welfare households are themselves more likely to become welfare recipients for longer durations. We also show that growing up in an ethnic environment characterized by welfare dependency has a significant effect on both the incidence and duration of welfare spells. About 80 percent of the difference in welfare participation rates between two ethnic groups in the parental generation is transmitted to the children.

George J. Borjas
Kennedy School of Government
Harvard University
79 John F. Kennedy Street
Cambridge, MA 02138
and NBER
gborjas@harvard.edu

Glenn T. Sueyoshi
Quantitative MicroSoftware
4521 Campus Drive
Irvine, CA 92612
glenn@eviews.com

ETHNICITY AND THE INTERGENERATIONAL TRANSMISSION OF WELFARE DEPENDENCY

George J. Borjas and Glenn T. Sueyoshi*

I. Introduction

The study of differences in socioeconomic status among ethnic groups in the United States is one of the core topics of social science.[1] Much of the research on ethnicity prior to the 1960s stressed the concept of a "melting pot" as a mechanism for the assimilation and integration of very diverse ethnic groups into the American mainstream. More recent research, however, emphasizes the ethnic differences that exist and persist along many dimensions of social, cultural, and economic background (Jiobu, 1990; Lieberson and Waters, 1988).

This paper investigates the relationship between ethnicity and welfare dependency. The analysis yields two key facts: There are sizable differences in the incidence of welfare receipt and in the duration of welfare spells across ethnic groups, and these differences tend to persist across generations. Although there has been extensive research on questions related to welfare dependency and the intergenerational transmission of this dependency in recent years, the existing literature ignores the important link between ethnicity and welfare (see Antel, 1992; Duncan, Hill, and Hoffman, 1988; Ellwood, 1989; Gottschalk, 1990; Moffitt, 1990; and Solon, et al, 1988).

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[1]For a sampling of very diverse methodological and ideological approaches to the study of ethnic differences, see the work of Alba (1990), Glazer and Moynihan (1963), Jiobu (1990), Perlmann (1988), Sowell (1981), and Steinberg (1989).

In view of the growing demographic importance of immigration, the speed of social and economic mobility in immigrant households has vital policy implications. There are large differences in skills, socioeconomic status, and welfare recipiency rates among the national origin groups that make up the immigrant flow (Borjas, 1990; Borjas and Trejo, 1991). If these differences are intergenerationally transmitted to U.S.-born ethnic groups, there might be significant long-run costs associated with policies which ignore the skill level of applicants in awarding entry visas.

The traditional model of intergenerational mobility in economics focuses on the role of parental investments in determining the human capital stock of their children (Becker, 1981). In contrast, the sociology literature stresses the importance of neighborhood or environmental variables (Coleman, 1988; Wilson, 1987). The operational hypothesis of the analysis presented in this paper is that ethnicity matters because it influences the "quality" of the environment in which human capital investments are made.[2] For given levels of parental inputs, persons who grow up in "high-quality" ethnic environments are likely to be exposed to different experiences (e.g., role models, social networks, job contacts, attitudes towards work and welfare dependency) than persons who grow up in a less advantageous ethnic milieu.[3] As a result, ethnicity is likely to play a crucial role in the intergenerational transmission process. Our analysis of the National Longitudinal Surveys of Youth (NLSY) reveals that persons who grow up in "welfare-prone"

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[2]The hypothesis that ethnicity may play an external effect on the human capital accumulation process was first formally developed in Borjas (1992).

[3]The analysis is obviously related to the literature on "neighborhood effects"; see Borjas (1995) and Case and Katz (1991) for recent examples. Jencks and Meyer (1990) provide an excellent survey of this literature, while Manski (1991) provides a critical appraisal.

ethnic environments are more likely to have both a higher incidence of welfare recipiency and longer-lasting welfare spells.

II. Ethnicity as a Human Capital Externality

The individual's decision to go on welfare is based on a comparison of the income opportunities available through the welfare system, W, and the income opportunities available through the labor market, Y. We interpret the benefit level W as net of any psychic costs (or "stigma") associated with being a welfare recipient. Define the index function:

(1) $$I = W - Y,$$

The person participates in the welfare system if $I > 0$. We are interested in determining the extent to which ethnicity *per se* affects the welfare participation decision.

Consider initially how ethnicity influences the worker's labor market opportunities. Assume that workers do not invest in their own human capital, so that the human capital stock of workers in generation $t+1$ is completely determined by the actions of generation t. The one-parent household has a utility function defined over the human capital stock of the child, k_{t+1}, and own consumption C_t:

(2) $$U = U(k_{t+1}, C_t).$$

The parent can either sell his human capital to the marketplace (at price R) or devote a fraction s_t of his time to the production of the child's human capital. Setting the price of C_t to unity gives the parent's budget constraint:

$$(3) \qquad R\,(1 - s_t)\,k_t = C_t\,.$$

Ethnicity is introduced by assuming that the average human capital stock of the ethnic group, $\bar{k}_t$, or "ethnic capital", has a spillover effect on the production of the human capital of children. The production function for child quality is given by:

$$(4) \qquad k_{t+1} = f(s_t k_t,\ \bar{k}_t)$$

Equation (4) indicates that two factors determine the human capital of children: (1) parental inputs as measured by $s_t k_t$, the "effective" amount of the parental human capital stock that goes into producing children's skills; and (2) the average human capital stock of the ethnic group in the parent's generation.

The hypothesis that ethnicity has external effects on human capital accumulation has strong roots both in sociology and in economics. For instance, Coleman (1988) stresses that the culture in which the individual is raised (which he calls "social capital") can be thought of as a form of human capital common to all members of that group. He argues that social capital alters the opportunity set of workers and has significant effects on behavior, human capital formation, and labor market outcomes.[4] Similarly, the influential work of Lucas (1988) uses an aggregate production function similar to (4) to analyze the process of economic development, and to "explain" why some countries may remain poor, while others grow richer (see also Romer, 1986; Barro, 1989).

The maximization of (2) subject to the budget constraint and the human capital production function generates a reduced-form equation determining the human capital stock of children:

$$(5) \qquad k_{t+1} = g(k_t,\ \bar{k}_t).$$

It is easy to show that $\partial k_{t+1}/\partial k_t > 0$ and $\partial k_{t+1}/\partial \bar{k}_t > 0$, so that there is a positive correlation between the human capital of children and parents, as well as between the human capital of children and the ethnic average.[5] The main implication of the ethnic capital model is that if the ethnic externality is sufficiently strong, skill differentials observed among ethnic groups are likely to persist for many generations and may never disappear. Put differently, exposure to an advantageous ethnic environment retards the downward slide of relatively skilled workers towards the mean; while exposure to a disadvantaged ethnic environment prevents the typical child in a relatively unskilled ethnic group from rising "too fast" towards the population mean.

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[4]Similarly, in his influential study of the underclass, Wilson (1986) argues that the presence of mainstream role models in poor neighborhoods serves an important social and economic function. In effect, the sociological literature argues that "neighborhood" effects (whether measured in terms of physical proximity to role models or as shared cultural values and attitudes) introduce important externalities in the human capital accumulation process.

[5]Apart from the usual concavity and differentiability restrictions on the utility and production functions, all that is required to prove these results is that child quality and consumption are normal goods in the utility function, and that parental human capital and ethnic capital are normal inputs in the production function.

In addition to its effect on human capital accumulation, ethnicity may also affect the net welfare benefit level, W. A simple application of the ethnic capital hypothesis suggests that both the probability that the parents receive welfare, p_t, as well as the welfare recipiency rate of the ethnic group in the parent's generation, $\bar{p}_t$, are likely to influence W. This "production function" for generating welfare benefits suggests that if a child is frequently exposed to welfare programs as an income-generating activity, the child may well have more information on the welfare opportunity set and may not suffer much stigma on becoming a welfare recipient. The ethnic capital hypothesis thus suggests that the index function in (1) can be rewritten as:

$$(6) \qquad I_{t+1} = W(p_t, \bar{p}_t) - Y(k_t, \bar{k}_t).$$

The welfare dependency of persons in the current generation, therefore, depends on their ethnic background because the average skills of the ethnic group, as well as the average welfare propensity of the group have spillover effects which influence the intergenerational transmission of "welfare proneness".

III. Data

The empirical study uses the 1979-1989 waves of the National Longitudinal Surveys of Youth. Respondents in the NLSY are aged 14-22 at the time of the initial survey in 1979. The person's ethnic background is determined from responses to the question: "What is your origin or descent?" Although most persons in the NLSY gave only one response to the question, about one third of the respondents gave multiple answers. In these cases, we use the main ethnic background (as identified by the respondent) to classify people into ethnic categories. Persons who did not provide a valid ethnic origin (or who have missing data for the other variables used in the analysis) are omitted from the study.

We analyze both the incidence of welfare recipiency and the duration of welfare spells. Our incidence measure is calculated using the entire NLSY panel. A person is defined to be a welfare recipient if he or she received income from the Aid to Families with Dependent Children (AFDC) program at any time during the 12-year span of the data. Our duration measure is given by the length of the *initial* spell of AFDC recipiency that took place between 1978 and 1989. We conducted a parallel analysis in which we defined welfare incidence in terms of the individual receiving AFDC income in the 1989 calendar year. The results were quite similar to those reported below (except for standard errors which were higher, but not enough to change conclusions for the key variables in the analysis).

To estimate the empirical counterpart of (6), we require data on both parental background and on characteristics of the ethnic environment where the NLSY respondents were raised. Information on whether the parents received welfare is available for the 79.6 percent of the NLSY children who resided with either parent during the initial survey year (1979).[6] We limit most of our analysis to this subsample of persons. There are obvious selection problems with this sample selection rule; it may be, for instance, that the most successful children moved out of the parental household earlier than other children. We show below, however, that this bias is minimal and does not greatly influence our results.

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[6]The parents of these NLSY respondents were asked about the sources of household income. A parental household is defined to be on welfare if the parents received public assistance in the previous calendar year.

It is worth noting that the raw data indicate a strong correlation between the parents' and the child's welfare recipiency: For the 6,210 persons in the NLSY subsample described above, the welfare participation rate of respondents raised in welfare households is 35.0 percent, while that of children raised in non-welfare households is 10.1 percent.[7] To isolate the impact of parental welfare recipiency (as opposed to parental skills) on the child's recipiency probability, we will control for two additional measures of parental background: the father's educational attainment and a dummy variable indicating if either parent was working at the time the children were 14 years old.

Additionally, we employ monthly data on AFDC usage between 1979 and 1989 to construct retrospective measures of the duration of welfare spells. The welfare durations used in our analysis are based upon responses to the annual question of whether the respondent or the spouse received AFDC income in each of the twelve months of the calendar year, cross-checking the monthly data against the annual reports on AFDC income. We define the onset of a spell of AFDC receipt as occurring when a person changes from non-recipiency to recipiency status. We then increment the spell length by one for each subsequent and consecutive month of reported receipt. The spell of AFDC is terminated in one of two ways: "normally," if the person stops receiving AFDC income, and "right-censored", if the welfare status or other essential data in the relevant calendar month is not reported for some reason. While we are not, in general, able to

[7] For our incidence analysis we employ 6,210 observations after all sample restrictions. Out of the original 12,686 observations, 2,221 were deleted because they were not interviewed in the 1989 wave; 341 were deleted because their place of residence at age 14 was either outside the United States or could not be ascertained; 1,469 were deleted because they did not report the father's education; 63 were deleted because their own education attainment was not ascertainable; 677 were deleted because they did not report an ethnic background; and 14 were deleted because of invalid codes in the welfare variable. Of the remaining 7,901 observations, 1,611 did not reside with their parents in 1979, and 80 were in ethnic groups that were eliminated from the sample for various reasons.

identify the cause of right censoring, among the important causes are non-interviews, refusal to respond to the AFDC receipt question, and missing data for relevant questions.[8] We employed this algorithm to construct a monthly welfare history spanning the 12 years from 1978 to 1989 for every NLSY respondent. Our empirical analysis focuses on the initial reported AFDC spell.

Our hypothesis stresses that the welfare dependency of NLSY respondents depend not only on parental influences, but also on the ethnic environment. We use the 1/100 1980 U.S. Census to calculate the welfare participation rate for each of the ethnic groups in the parents' generation.[9] The Census data report the ancestral background of U.S.-born residents (obtained from questions resembling the self-reported ethnic background in the NLSY). To increase the probability that the estimated welfare participation rates represent the welfare propensity of the ethnic milieu in which the NLSY respondents were raised, we restrict the 1980 Census sample to household heads aged 35-64. We also use the Census data to calculate two additional measures of the average skills of the ethnic group: the average educational attainment of men in the group and the group's male labor force participation rate.[10]

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[8] It is worth emphasizing that under the maintained hypothesis that missing data occur randomly, we artificially censor durations when important explanatory variables (e.g., education, SMSA of residence, marital status) are not reported in a given period. Thus, if data on current residence are not available in the eighth month, we end the AFDC spell at seven months and note that the actual duration is at least that length. In so censoring the data, a number of durations are lost because they are missing information for the first period.

[9]The ethnic characteristics are calculated using a 20 percent random sample of the 5/100 A File of the 1980 Public Use Sample. We also constructed comparable ethnic characteristics from within the NLSY itself. Although the results do not depend on which of the two sets of aggregate variables we use, the discussion below only reports the regressions that use the Census data (which are calculated over much larger samples and hence are likely to contain much less sampling error).

[10] We also constructed the "ethnic capital" variables by using geographic variation within each ethnic group (and obtained very similar results to those reported below). In particular, we defined the mean probability of welfare recipiency (or educational attainment, or labor force participation) for members of the ethnic group that lived in state k in 1980. This more detailed specification of the measure of ethnic capital should, in principle, allow for a more complete analysis of the ethnic capital model since external effects are more likely to be significant if the NLSY respondent actually had physical contact with members of the ethnic group. However, the empirical

The second column of Table 1 documents the dispersion in welfare participation rates within our NLSY sample. For example, respondents of English, French, German, and Irish origin have welfare participation rates hovering in the neighborhood of 8 percent, while 16 percent of Mexicans, 23 percent of blacks, and 18 percent of Puerto Ricans receive AFDC. Table 1 also reports that the approximate average spell of welfare receipt lasted about 18 months for respondents of Mexican, Italian, and Portuguese origin, but lasted 27 months for blacks and Puerto Ricans. Finally, the recipiency rates estimated in the Census for the parental generation, and reported in the last column of the table, reveal equally large ethnic differences in the fraction of the group receiving welfare during a calendar year. Households of English or German origin had welfare participation rates of 4 to 5 percent, while 12.5 percent of Mexicans, 22 percent of blacks, and 31 percent of Puerto Ricans report welfare receipt.

IV. Ethnicity and the Incidence of Welfare Receipt

Let $Z_{ij}(t)$ be an unobserved latent variable measuring the incentives of individual i in ethnic group j in generation t to enter the welfare system. The statistical model that separately identifies the impact of parental and of ethnic capital as determinants of welfare participation behavior in generation t is given by:

$$(7) \qquad Z_{ij}(t) = \gamma_0 + \gamma_1 \, x_{ij}(t\text{-}1) + \gamma_2 \, \bar{x}_j(t\text{-}1) + \varepsilon_{ij}(t)$$

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evidence in Borjas (1995) suggests that the correct geographic level to measure these external effects is the *neighborhood* where the child was raised, not the state. Unfortunately, we cannot link the geographic data available in the public use samples from the Census to neighborhood data in the NLSY.

where $x_{ij}(t\text{-}1)$ is a vector of relevant skill characteristics of the parent; and $\bar{x}_j(t\text{-}1)$ gives the average characteristics for the ethnic group in the parent's generation.

Although $Z_{ij}(t)$ is not observed, the data report an indicator variable $p_{ij}(t)$ which takes on the value of unity if the individual receives public assistance and zero otherwise. The variable $p_{ij}(t)$ is defined by:

$$(8) \qquad p_{ij}(t)\begin{cases} = 0, & \text{if } Z_{ij}(t) < 0, \\ = 1, & \text{otherwise.} \end{cases}$$

To separately estimate the impact of parental and ethnic welfare recipiency on welfare propensities, we employ a two-stage estimator for a random effects probit. The stochastic structure of this model is best understood by providing an alternative derivation of equation (7). The welfare participation of generation t depends both on parental characteristics and on a vector of ethnic fixed effects:

$$(9) \qquad Z_{ij}(t) = \gamma_1\, x_{ij}(t\text{-}1) + d_j + u_{ij}(t),$$

where d_j is the ethnic fixed effect, and $u_{ij}(t)$ is assumed to be *iid* normal so that a model based upon (9) may be estimated using maximum likelihood (ML) probit. The hypothesis that ethnicity acts as a human capital externality implies that the fixed effects can be partly explained by differences in the "quality" of the ethnic environment. Let $\bar{x}_j(t\text{-}1)$ measure relevant aspects of the ethnic environment (such as the welfare participation rate of the ethnic group in the parental generation). The variable d_j can then be given the structural specification:

(10) $$d_j = \gamma_0 + \gamma_2\, \bar{x}_j(t\text{-}1) + v_j.$$

Substituting (10) into (9) yields equation (7). Note, however, that the disturbance in equation (7) is given by $\varepsilon_{ij}(t) = v_j + u_{ij}(t)$. Because it has an ethnic group component, the error term has the stochastic structure of a probit random effects model (Heckman and Willis, 1977), where the errors are correlated for observations belonging to the same ethnic group. Estimation procedures that ignore this correlation provide badly biased standard errors (particularly for the coefficients of the group-level variables).[11]

The alternative derivation of (7) suggests a computationally simple two-stage estimator. In the first stage, we apply ML probit to (9) and obtain the coefficients of the individual-level variables and the ethnic fixed effects. In the second stage, we estimate the parameters in (10) by regressing the fixed effects on the ethnic-specific characteristics using generalized least squares to correct for estimation variance. Borjas and Sueyoshi (1994) show that this two-stage procedure for estimating probit models with structural group effects is asymptotically unbiased as group sizes grow, and consistent when both group sizes and the number of groups increase.[12]

We begin our analysis with a first-stage probit specification which holds constant only a minimal set of demographic variables (gender, age, and a dummy variable indicating if the

[11]The importance of this bias for hypothesis testing in linear models containing group-level variables was noted by Moulton (1986).

[12]Borjas and Sueyoshi (1994) also report Monte Carlo showing that the two-stage estimator substantially outperforms the computationally more complex one-stage random effects probit when there are large group sizes. Sueyoshi (1992) provides theoretical results for the rates at which the number of groups must grow relative to the sizes of the groups for the consistency results to hold.

NLSY respondent is foreign-born, indicator variables for Census region of residence, two indicator variables for high and low unemployment rates in the current area of residence, and a dummy variable indicating if the NLSY respondent lived in a two-parent household at age 14).[13] By restricting the controls to these "exogenous" factors, the estimated intergenerational transmission parameters do not net out the impact of parental and ethnic background on other socioeconomic variables, particularly the NLSY respondent's educational attainment.[14] It will be seen below that many of the results are not affected substantially when the fixed-effect probits also control for the child's educational attainment.

The first line of Table 2 presents parameter estimates and asymptotic standard errors for relevant coefficients of the simplest specification, where the only variable in the parental vector indicates if the parents received welfare in 1979, and the only variable in the ethnic capital vector gives the mean welfare recipiency of the ethnic group in the parental generation.[15] Both coefficients are statistically significant, with the ethnic group welfare coefficient roughly 4 times as large as the parental welfare receipt coefficient. To provide some perspective on these results, the parameter estimates indicate that the children of welfare recipients have a welfare

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[13]The regression includes a small number of foreign-born persons (5 percent of the sample). Our sample, however, is restricted to persons who are residing in the United States at age 14. We also conducted the analysis on the smaller sample of U.S.-born persons, without any appreciable change in the results.

[14] Many studies of welfare recipiency also make sample restrictions based on the marital status of the householder. One can argue, however, that marital status--particularly in the context of welfare recipiency--is an endogenous variable. After all, persons can easily switch among the various matrimonial states (with little change in *actual* household composition) in order to meet the eligibility rules of the welfare program. To test the sensitivity of our results, we conducted parallel analyses of the key specifications that included variables measuring the marital status of the NLSY respondent or that restricted the sample to particular groups (such as those married, spouse absent). The results provided by these alternative specifications are similar to those reported in this paper.

[15] The parental coefficients are derived from the first stage estimation; the ethnic capital coefficients are estimated in the second stage regression.

participation rate that is 16.9 percentage points higher than the rate for children of non-recipients.[16] The regressions also show that even after controlling for parental welfare participation, a one percentage point increase in the recipiency rate of the ethnic group in the parents' generation increases the participation rate of the child by .67 percentage points.[17]

There is, therefore, a very strong link between the recipiency rate of the ethnic group across generations resulting from both parental background and ethnic externalities. A summary statistic measuring this relationship is given by the sum of the coefficients $\hat{\gamma}_1 + \hat{\gamma}_2 = .770 + 3.065 = 3.835$. Abstracting from the constant term, $(\gamma_1 + \gamma_2) \cdot \phi(x'\beta)$ gives a predicted derivative of .8394 for the offspring of the average father in ethnic group j.[18] Thus, the results suggest that if two ethnic groups are 10 percentage points apart in welfare recipiency in generation t, they will be 8.4 percentage points apart in generation t+1. It is important to note that this linkage is much stronger than is suggested by studies which ignore the ethnic externality and simply correlate the welfare participation rate of parents and children. For instance, the probit estimation of (7) omitting the ethnic variable yields:

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[16] To convert the probit coefficients into units that estimate $\partial p/\partial x$, the probit coefficient β must be multiplied by $\phi(x'\beta)$, where ϕ is the density function for the standard normal. We approximate this term by $\phi(z)$, where z is the quantile associated with the mean welfare participation probability. See the notes to the table for additional details.

[17] The R^2 of the second-stage equation is about .4, so that the ethnic welfare propensity "explains" a little under half of the variation across groups in the standardized welfare participation rates.

[18] The thought experiment predicts the participation probability of the child of the average father in a particular ethnic group. The parental welfare participation probability is then the same as the measure of the ethnic capital variable.

(11) $$Y_{ij}(t) = Z\pi + .841\, p_{ij}(t\text{-}1),$$

where Z is the same vector of socioeconomic characteristics held constant in Table 2, and the standard error of the parental coefficient is .053. Equation (11) implies that a 1 percentage point increase in the welfare participation rate of parents increases the participation of NLSY children by only 0.184 percentage points (.841 × .219). Thus, the omission of ethnic externalities greatly underestimates (by about three-fourths), the extent to which the welfare participation rate of ethnic groups is linked across generations.

It is important to determine if the strong estimated relationship between the welfare recipiency of the group in the t-th generation and the recipiency rate in the parental generation arises because the welfare participation rate of the group is only a proxy for the group's skills, or whether there is a separate impact of group welfare recipiency, even after controlling for the group skill level. The second row of Table 2 indicates that the relationship between the NLSY respondent's welfare receipt probability and the mean recipiency of the parental group remains, even if other characteristics of the parents and the groups are controlled for. In particular, we expand the first stage estimates to include two additional measures of parental skills: the father's educational attainment, and a dummy variable indicating whether either parent worked when the NLSY respondent was 14 years old. These variables, as expected, are negatively correlated with welfare participation of the children (although the employment variable is statistically insignificant). Note, however, that externalities remain--adding together the parental and ethnic welfare coefficients indicates that a one percentage point increase in the recipiency rate of the parental generation increases the recipiency rate of the children's generation by 0.73 percentage points.

It is of interest to determine if the impact of the parental and ethnic welfare recipiency variables remains when we control for the respondent's skills. We reestimated the models above including a variable indicating the respondent's educational attainment. The third row of Table 2 shows that adding own educational attainment reduces slightly the impact of both parental and ethnic welfare recipiency, but these variables still play a role in intergenerational transmission. In particular, parental welfare receipt is associated with a 13.6 percentage point higher probability of child welfare receipt, while a 1 percentage increase in the ethnic welfare participation rate increases child recipiency by 0.44 percentage points. Thus, both parental characteristics and ethnic externalities influence the welfare participation behavior of children above and beyond the traditional transmission mechanism that works through educational investments.

The second panel of Table 2 documents similar results when the analysis is restricted to a subsample of 3,021 female NLSY respondents.[19] The results in the second panel imply that parental and ethnic welfare participation have, if anything, a stronger impact upon the participation rates of NLSY daughters than on the sample as a whole. Focusing on the most general specification (row 3), we find that parental welfare recipiency increases the probability of welfare receipt by 23 percentage points, while increasing the ethnic welfare recipiency rate by 1 percentage point increases child welfare use by 0.94 percentage points. Adding the coefficients together implies that a one percentage point increase in the ethnic recipiency rate yields a 1.17

[19] Of these individuals, 663 respondents report welfare receipt (663/3,021 = .22).

percentage point increase in the welfare propensity of the youth. Put differently, the ethnic externality on welfare recipiency is so strong that women raised in ethnic environments with high welfare participation rates seem to have little chance of escaping a lifestyle of dependency.

As noted above, the subsample of NLSY children who lived with their parents in 1979 is not randomly selected. The bottom panel of Table 2 reports the estimated coefficients when persons who were not living with their parents in 1979 are added to the sample, and parental welfare use is dropped as an explanatory variable so that the parental background vector contains only two variables, the father's educational attainment and a dummy variable indicating whether either parent was working at the time the respondents were 14 years old.[20] It is evident that the children of highly-educated or working parents have significantly lower welfare propensities. For example, (again focusing on specification (3)), an additional year of father's education reduces the welfare participation probability of the child by 1.1 percentage points. Similarly, persons who grew up in a household where at least one parent worked have a participation rate that is 7.4 percentage points lower than persons who grew up in non-working households.

Even after controlling for these measures of parental background, the analysis indicates that the impact of the welfare participation rate of the ethnic group in the parent's generation is similar to the estimates derived from the subsamples described above. In particular, the intergenerational coefficient is 2.552. This coefficient implies that a 1 percentage point increase in the ethnic welfare recipiency rate increases the welfare probability by 0.606 percentage points. To the extent that the parental variables proxy for the unobserved parental welfare participation,

[20] This full sample contains 7,834 respondents, 1,207 of whom report welfare receipt (1,207/7,834 = 0.15).

the coefficient of the ethnic welfare participation rate in the bottom panel of Table 2 should resemble the corresponding coefficient estimates in the upper panel. This appears to be the case. so that the bias introduced by restricting the analysis to NLSY children who resided with their parents in 1979 is minimal.

Before proceeding to an analysis of the duration of welfare spells, it is instructive to illustrate the sensitivity of our results to a key alternative specification of the model. The summary data reported in Table 1 shows that some ethnic groups have very high welfare participation rates (particularly blacks and some of the Hispanic groups), so that some of our results might be driven by the presence of one or two "outlying" ethnic groups. It is important, therefore, to determine if the results are sensitive to the omission of these outlying groups from the study.

Figure 1 provides a graphical summary of the second-stage estimates revealing the positive relationship between the group ethnic welfare variable and the estimated group effects from the first-stage probit estimation. The figure strongly suggests that the results are not an artifact generated by outliers generated by the high welfare groups (Puerto Rican, black). This insight is confirmed by the more formal statistical analysis reported in Table 3. This table shows that the results remain essentially unchanged if we restrict the analysis to non-Hispanic groups. Adding the coefficients in the simplest specification (row 1) reveals that a one percentage point increase in the recipiency rate of the parental generation raises the participation rate of the group in the children's generation by 0.91 percentage points. The regressions also show that the coefficients in the simpler specification remain stable when we further restrict the sample to white, non-Hispanic ethnic groups. A one percentage point increase in the ethnic recipiency rate of the parental generation then raises the rate of the group in the children's generation by 0.61

percentage points. Note, however, that the coefficient of the parental recipiency rate becomes statistically insignificant when the regression model also includes additional variables measuring the educational attainment and labor force participation status of both the parent and the group.

V. Ethnicity and the Duration of Welfare Receipt

Let T be a continuous, non-negative random variable representing the duration of a spell of welfare receipt. We use standard techniques for the estimation of duration models to describe the influence of parental and ethnic welfare participation on T, in an analysis that parallels the incidence discussion above. As noted earlier (see Table 1), our data indicate sizable ethnic differences in the duration of welfare spells. We are primarily interested in the question of whether there is evidence of intergenerational transmission of welfare dependency. Because we do not know the duration of parental welfare spells, we focus on the relationship between the spell length of children and the welfare incidence of parents; the descriptive statistics reported earlier suggest that there is a positive correlation between these variables.

Our econometric specification is a variant of the proportional hazards model (Cox , 1972) which is adapted to handle the monthly frequency of our observations on duration (Prentice and Gloeckler, 1978), and our two-stage techniques for analyzing group effects. While more extensive descriptions of the techniques are available in the literature, we provide a brief description of the approach.

Let $\lambda(t) = \lim_{t \to 0} \Pr(t \leq T \leq t+dt \mid T \geq t)/dt$ be the hazard rate associated with T. We wish to model the probability of observing T in a discrete interval $(t_l, t_u]$ as functions of observable characteristics that enter the hazard through the variables X and parameters β.[21] A general specification for the hazard rate may be denoted $\lambda(t, X, \beta)$. From standard results , the survivor function $S(t, X, \beta) = \Pr(T \geq t) = \exp(-\exp(\int_0^t \lambda(s, X, \beta)\, ds))$, and the density function for durations $f(t, X, \beta) = \lambda(t, X, \beta)S(t, X, \beta)$ may be expressed in terms of our parameterization for λ. For interval duration data, standard likelihood contributions may be based upon the survivor function $S(t, X, \beta)$. An observed duration in the interval $(t_l, t_u]$ then contributes terms of the form $S(t_l, X, \beta) - S(t_u, X, \beta)$ to the likelihood function. Observations that are right censored at t_l contribute survivor terms of the form $S(t_l, X, \beta)$.

An alternative parameterization of the likelihood is possible using conditional survivor functions. Define the conditional survivor for $(t\text{-}1, t]$ as $\alpha(t, X, \beta) = \Pr(T > t \mid T > t\text{-}1) = S(t, X, \beta)/S(t\text{-}1, X, \beta)$. We can express the survivor in terms of the set of α's: $\prod_{j-1}^{t} \alpha(j, X, \beta)$. A likelihood function can then be written in a familiar "probit"-style form by including a binary outcome for each individual-period combination associated with a given $\alpha(j, X, \beta)$. Thus, if an individual is observed with a welfare spell of 9 months, we define 9 observations on a (0, 1) indicator for whether the individual remained on welfare after the interval, with the first 8 indicators coded as 1's and the 9th coded as a 0. The probability of a 1 in a given interval is given by the corresponding $\alpha(j, X, \beta)$ for the period.

To complete the specification of the model, we must choose the conditional survivor functions α. While one could choose a simple probit specification for each period, we present

[21] Note that both our normal and right censored observations on welfare durations are handled naturally in this interval framework; the former by letting t_l and t_u bound the observed monthly duration, the latter case by letting $t_u \rightarrow \infty$.

results for a flexible baseline, proportional hazard model: $\alpha(s, X(s), \eta, \beta) = \exp(-\exp(\eta_s + X(s)\beta))$, where the $X(s)$ are the explanatory variables observed in month s; β is the corresponding vector of parameters; and η are a set of period-specific constants.[22] It is worth noting that this parameterization of the proportional hazards model has an interpretation that is similar to the probit specification above. We may define for each period s a period-specific analogue to (7) and (10) which contains a period dummy,

$$(12) \qquad Z_{ijs}(t) = d_j + \eta_s + \gamma_1 X_{ijs}(t-1) + u_{ijs}(t),$$

where u is distributed as an *iid* Type-I extreme value random variable.[23]

Table 4 summarizes the results of our analysis by reporting the relevant coefficients from the first and second stages of the estimation process, and Figure 2 provides a graphical summary of the second-stage estimates showing the negative relationship between the parental ethnic group welfare propensity and the hazard of exit from welfare use. It is evident that both parental welfare recipiency and the welfare recipiency rates of the parental group have an important effect on the duration of welfare spells in the NLSY, with both sets of variables lowering the hazard rate. Even when additional controls for parental skills, additional variables describing the skills of the group, and own educational attainment of the NLSY respondent are introduced into the

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[22] It is easy to show that for this specification, the η may be interpreted as integrated baseline hazards. See Kiefer (1988) and Sueyoshi (1995) for further details on this parameterization of the standard proportional hazards model.

[23] Due to concerns that our results might be sensitive to the presence of unobserved heterogeneity or other misspecification of the functional form, we experimented with other, non-proportional hazard specifications involving different assumptions about this error term. Our results did not differ substantively, so we present only the proportional hazards results.

model, both welfare variables are negative and statistically significant at conventional levels of significance.

In order to provide a more meaningful interpretation of the coefficients of the duration model, we calculate the average spell length (in months) for various values of the parental and group variables. The results of this simulation are presented in Table 5 for the three specifications, and all three samples corresponding to Table 4.[24] The table presents both the predicted censored mean (where predictions lasting more than 52 months are censored at 52 months), as well as the probability that the welfare spell will last more than 52 months. For example, focusing on the specification (1) of the sample of persons who lived with their parents, the children of parents who are not welfare recipients will have a predicted average censored duration of 17.35 months of welfare receipt, with 9.6 percent of those experiencing spells that are longer than 52 months. In contrast, the children of welfare recipients will have an average censored duration of 22.19 months, and a 17.7 percent probability of having a spell lasting longer than 52 months.

To illustrate the magnitude of the impact of the group's recipiency rate on the duration of the children's spells, we shift the mean group recipiency probability by 10 percentage points. Again focusing on the first set of results, we find that the mean duration of welfare receipt

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[24] The simulations assume an individual with average characteristics for the other variables. We compute the mean durations and censoring frequency by selecting a value for the relevant welfare variable, and computing the predicted probabilities of exiting from welfare at each duration. A lower bound estimate of the expected duration is given by $E(T^*) = \sum_{j=1}^{52} j \cdot P(T=j) + 52 \cdot P(T>52)$. The parental welfare variable is evaluated at 0, representing parents who do not receive AFDC, and 1, representing recipients. The ethnic welfare variable is evaluated at the actual overall ethnic group propensity for the sample, as well as for the latter value increased by 10 percentage points.

increases from 19.1 to 22.2 months, and the probability that the spell lasts more than 52 months increases from 12.3 to 17.8 percent. Similar results are observed for the alternative specifications and samples. with mean durations increasing by roughly 6 months, and the frequency of right-censored observations approximately doubling. Thus, the empirical evidence indicates a strong intergenerational relationship in welfare dependency (in terms of the duration of welfare spells) resulting from both parental and ethnic influences.

VI. Measurement Error

Because the empirical evidence summarized in the last two sections is so striking, it is worth examining if a spurious correlation between the average welfare propensity of the ethnic group in generation t and the welfare propensity of children in generation t+1 may be driving the results. Such a spurious correlation is introduced by measurement error in parental welfare participation.

The analysis of measurement error is complicated by two factors. First, the independent variable measured with error is a binary response variable, and this imposes a particular stochastic structure on the measurement error. Second, the models estimated in the previous section are highly nonlinear, and the typical formulae derived for measurement error in linear models do not apply (see Yatchew and Griliches, 1985). To isolate the problem introduced by measurement error in parental welfare recipiency, we abstract from the nonlinear nature of the models and derive some results for a linear probability model in which the parental welfare recipiency and the child's welfare recipiency are linked as in:

$$(13) \qquad p_t = \alpha + \delta\, p_{t\text{-}1} + \varepsilon,$$

where p_t is a dummy variable indicating if the t-th generation is receiving welfare. For simplicity, we omit subscripts indicating individuals and ethnic groups. Note that the true model does not contain an ethnic effect.

Let $x_{t\text{-}1}$ be the observed parental welfare participation dummy, where $x_{t\text{-}1} = p_{t\text{-}1} + \omega$. If $p_{t\text{-}1} = 1$, then $\omega = -1$ with probability r_w (the fraction of welfare recipients who incorrectly provide their recipiency status), and $\omega = 0$ with probability $1 - r_w$. Similarly, if $p_{t\text{-}1} = 0$, then $\omega = 1$ with probability r_{nw} (the fraction of non-welfare recipients who claim to receive public assistance), and $\omega = 0$ with probability $1 - r_{nw}$. This implies that:

$$(14) \qquad E(\omega) = r_{nw} - (r_w + r_{nw})\, p_{t\text{-}1}$$

The measurement error does not have zero mean and is correlated with $p_{t\text{-}1}$. It is well known (Aigner, 1973) that if p_t is regressed on $x_{t\text{-}1}$, the probability limit of the coefficient is given by:

$$(15) \qquad \text{plim}\ \hat{\delta} = h\delta,$$

where $h = 1 - r_w - r_{nw}$.

Suppose that an additional measure of parental welfare participation is available, the welfare participation rate of the ethnic group in the parents' generation, $\bar{p}_{t-1}$. A generic form of the regression model we estimated earlier is:

(16) $$p_t = \alpha + \theta_1 x_{t-1} + \theta_2 \bar{p}_{t-1} + \varepsilon'.$$

Because the ethnic mean proxies for the imperfectly measured parental welfare participation, both parameters in the coefficient vector θ may be significant even if the true model is given by (13). Assume that the measurement error in x_{t-1} "washes out" when aggregating within the ethnic group so that the mean observed welfare participation rate in the ethnic group is measured without error. It can then be shown that:[25]

(17) $$\text{plim } \frac{\hat{\theta}_2}{\hat{\theta}_1} = \frac{1-h^2}{h\pi},$$

where π is the fraction of the variance in parental welfare participation rates that is within-group.

Equation (17) can be used to assess the importance of measurement error in the analysis. Over 95 percent of the variance in welfare participation rates is within-group, so that π is close to one. Although there are no available estimates of h (which is related to the fraction of households who correctly report their welfare recipiency status), there are analogous estimates of

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[25]The probability limits for the coefficients are as follows:

$$\text{plim } \hat{\theta}_1 = \frac{h\pi\delta}{1-h^2(1-\pi)},$$

$$\text{plim } \hat{\theta}_2 = \frac{(1-h^2)\delta}{1-h^2(1-\pi)}.$$

Note that the sum of the two probability limits does not add up to the true δ (which would be the case if the independent variable is not binary; see Borjas, 1992). The reason is that measurement error in binary variables is correlated with p_{t-1}, and hence the ethnic welfare participation rate is not a valid instrument for parental recipiency.

the number of persons who correctly report their union status. Freeman's (1984) systematic study of measurement error in reporting union status concludes that $h \approx .9$.[26] If response errors on welfare recipiency are of the same order of magnitude, the ratio in (14) would be only .21. In fact, the empirical results presented earlier indicate that $\hat{\theta}_2 / \hat{\theta}_1 > 1$. To explain these results in terms of measurement error would require that h is less than .5 or .6, so that over one-third of the observations in the sample incorrectly report their welfare status. This is unlikely because the welfare recipiency of the parental household was *not* obtained from retrospective responses provided by the NLSY children, but is ascertained from questions posed to the parents themselves regarding their income sources in 1979.

Of course, these measurement errors are not strictly correct because (17) is derived in a linear probability framework, while our estimates are derived from nonlinear ML specifications (probit, extreme value binary outcome). The extent of this misspecification, however, may not be very important. If the regression in the first row of Table 2 had been estimated using the linear probability model, the resulting coefficients (and standard errors) are $\hat{\theta}_1 = .1683$ (.0110), and $\hat{\theta}_2 = .3752$ (.0711). To dismiss these results because of measurement error would require a substantial misreporting of welfare participation in the parental population.

It would be useful to derive formulae equivalent to (17) for measurement errors in probit and ML duration models. However, measurement errors in these nonlinear models lead to misspecification of the underlying probability functions unless the reporting errors themselves follow restrictive forms. In the probit case, it is easy to show that the probability limit of the

[26]Poterba and Summers (1986) report measurement errors of approximately the same order of magnitude in their study of transitions in and out of unemployment.

ratio (17) is unchanged if the errors are normally distributed and the model is estimated using a probit. The measurement errors in parental welfare recipiency, however, follow a correlated Bernoulli process so that the disturbance in the reduced-form probit is a mixture of normal and binomial random variables. As long as the fraction of persons reporting parental welfare status is small, the resulting mixture may be closely approximated by a normal distribution, and the results presented in this section provide a reasonable representation of the true structure.

VII. Summary

This paper investigated the relationship between ethnicity and welfare dependency. The operational hypothesis is that ethnicity has external effects on the human capital accumulation process. This hypothesis implies that the welfare participation of the current generation depends not only on variables describing the skills and other relevant characteristics of the parents, but also on variables describing the skills and propensities towards work and welfare in the ethnic milieu.

The study of data drawn from various sources (including the National Longitudinal Surveys of Youth and decennial U.S. Censuses) indicated that ethnic differences in welfare propensities are large and persistent. As suggested by the conceptual framework, ethnic differences in welfare participation behavior are transmitted across generations for two distinct reasons. First, the children of welfare households are more likely to become welfare recipients. In addition, growing up in an ethnic environment characterized by welfare dependency has a direct effect on welfare participation in the next generation (after holding parental skills constant).

The documented link between ethnicity and the intergenerational transmission of welfare recipiency has provocative policy implications. As a result, some words of caution are in order. For instance, although the analysis shows that ethnicity matters, it does not say *why* it matters. Assessing the consequences of alternative policy proposals requires a much deeper understanding of the channels through which the transmission of ethnic fixed effects takes place.

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TABLE 1. SAMPLE STATISTICS FOR THE FREQUENCY AND DURATION OF WELFARE RECEIPT, BY ETHNICITY.

	Incidence		Duration			Parental Recipiency
	(1)	(2)	(3)	(4)	(5)	(6)
Ethnic Group/Code	Sample Size	Percent Receiving Welfare	Sample Size	Mean Duration (months)	Percent > 52 months	Percent Receiving Welfare
American Indian (AM-IN)	327	16.21	49	16.49	10.20	12.48
American (AMER)	356	12.64	43	24.05	13.95	4.06
Black (BLK)	1569	22.56	326	26.58	13.80	21.57
Cuban (CU)	65	4.62	3	29.33	33.33	7.58
English (ENG)	939	9.05	78	20.18	17.95	4.56
Filipino (FIL)	---	---	1	12.00	0.00	6.67
French (FR)	157	8.92	13	5.85	0.00	7.81
German (GER)	802	8.35	62	16.61	8.07	3.93
Greek (GRK)	15	13.33	2	9.00	0.00	4.78
Irish (IR)	512	9.38	44	12.89	2.27	5.76
Italian (ITAL)	301	6.31	18	18.50	11.11	4.98
Mexican (MEX)	641	15.91	86	18.31	10.47	12.48
Other Hispanic (OHS)	94	10.64	8	22.50	12.50	14.37
Pacific Islander (PI)	---	---	2	18.00	0.00	12.02
Polish (POL)	136	3.68	5	21.80	0.00	4.36
Portuguese (POR)	57	17.54	8	18.00	25.00	7.89
Puerto Rican (PRT)	148	17.57	26	26.65	11.54	31.22
Russian (RUS)	28	3.57	---	---	---	2.77
Scottish (SCT)	60	3.33	2	9.50	0.00	3.80
	6120	13.66	776	21.86	12.11	---

Notes: The code in parentheses gives the group's label in the figures presented below. All calculations are conducted in the sample of persons who lived with their parents in 1979 (at the time of the initial survey). The mean duration reported in this table is censored at 52 months (i.e., duration takes on a value of 52 if the spell lasted longer than 52 months). This computation does not adjust for other right-censored observations, those censored observations are treated as real failures for purposes of this computation.

TABLE 2. EFFECTS OF PARENTAL AND ETHNIC GROUP VARIABLES ON THE INCIDENCE OF WELFARE RECEIPT

	Parental Variables			Ethnic Group Variables			
Model	Welfare	Education	LFP	Welfare	Education	LFP	Own Education
A. Lived With Parents							
(1)	0.770	---	---	3.065	---	---	---
	(0.060)			(0.427)			
(2)	0.668	-0.049	-0.089	2.682	0.009	-2.490	---
	(0.062)	(0.007)	(0.074)	(1.154)	(0.044)	(1.814)	
(3)	0.611	-0.038	-0.086	2.030	0.021	-2.272	-.178
	(0.064)	(0.007)	(0.075)	(1.178)	(0.045)	(1.849)	(.022)
B. Female, Lived with Parents							
(1)	0.959	---	---	4.293	---	---	---
	(0.073)			(0.570)			
(2)	0.863	-0.055	-0.028	2.910	0.039	-3.143	---
	(0.077)	(0.008)	(0.092)	(1.511)	(0.056)	(2.360)	
(3)	0.797	-0.045	-0.006	3.199	0.050	-3.084	-.202
	(0.078)	(0.008)	(0.093)	(1.580)	(0.059)	(2.465)	(.028)
C. Full Sample							
(1)	---	---	---	4.088	---	---	---
				(0.363)			
(2)	---	-0.060	-0.330	2.168	0.022	-3.189	---
		(0.006)	(0.059)	(0.950)	(0.037)	(1.491)	
(3)	---	-0.044	-0.271	2.552	0.040	-2.972	-.213
		(0.006)	(0.061)	(1.034)	(0.040)	(1.617)	(.017)

Notes: Standard errors are reported in parentheses. In sample A, the mean welfare participation is .137, and $z = \Phi^{-1}(.137) = -1.096$ where Φ is the standard normal cumulative distribution function, so that the adjustment factor is $\phi(z) = .219$, where ϕ is the density function. The derivatives reported in the text are then derived by multiplying coefficient estimates by .219. The recipiency rate in sample B is .220, and the adjustment factor is .296. The recipiency rate in sample C is 0.154, and the adjustment factor is .237.

TABLE 3. EFFECTS OF PARENTAL AND ETHNIC GROUP VARIABLES IN SUBSAMPLES OF ETHNIC GROUPS

	Parental Variables			Ethnic Group Variables		
Model	Welfare	Education	LFP	Welfare	Education	LFP
A. Non-Hispanic Sample, Lived With Parents						
(1)	0.793	---	---	3.605	---	---
	(0.086)			(0.453)		
(2)	0.668	-0.055	-0.152	2.575	-0.102	0.871
	(0.069)	(0.012)	(0.082)	(0.638)	(0.043)	(1.537)
B. Non-Hispanic Female Sample, Lived with Parents[b]						
(1)	0.995	---	---	5.211	---	---
	(0.062)			(0.610)		
(2)	0.888	-0.060	-0.057	4.498	-0.079	1.077
	(0.042)	(0.014)	(0.103)	(0.724)	(0.050)	(1.950)
C. White Non-Hispanic Sample, Lived with Parents						
(1)	0.943	---	---	3.235	---	---
	(0.073)			(0.897)		
(2)	0.781	-0.073	-0.159	0.883	-0.161	1.046
	(0.085)	(0.005)	(0.136)	(1.281)	(0.038)	(1.275)
D. White Non-Hispanic Female Sample, Lived with Parents						
(1)	1.101	---	---	3.858	---	---
	(0.082)			(1.323)		
(2)	0.958	-0.082	-0.061	0.990	-0.203	1.340
	(0.074)	(0.008)	(0.183)	(2.002)	(0.047)	(1.403)

Notes: Standard errors are reported in parentheses. The recipiency rate in sample A is .132, and the adjustment factor is .206. The recipiency rate in sample B is .217, and the adjustment factor is .294. The recipiency rate in sample C is .077, and the adjustment factor is .145. The recipiency rate in sample D is .100, and the adjustment factor is .175.

TABLE 4. EFFECTS OF PARENTAL AND ETHNIC GROUP VARIABLES ON THE HAZARD OF EXITING FROM A SPELL OF WELFARE RECEIPT

	Parental Variables			Ethnic Group Variables			Own
Model	Welfare	Education	LFP	Welfare	Education	LFP	Education
A. Lived With Parents							
(1)	-0.305	---	---	-1.943	---	---	---
	(0.091)			(0.734)			
(2)	-0.290	0.025	-0.058	-3.838	-0.052	-3.277	---
	(0.094)	(0.012)	(0.112)	(1.992)	(0.075)	(3.190)	
(3)	-0.280	0.023	-0.062	-4.132	-0.059	-3.575	0.024
	(0.032)	(0.013)	(0.112)	(1.980)	(0.074)	(3.146)	(0.029)
B. Female, Lived with Parents							
(1)	-0.305	---	---	-2.707	---	---	---
	(0.102)			(0.819)			
(2)	-0.294	0.036	-0.115	-4.924	-0.031	-4.211	---
	(0.105)	(0.014)	(0.131)	(2.314)	(0.085)	(3.619)	
(3)	-0.293	0.036	-0.116	-4.989	-0.032	-4.283	0.002
	(0.106)	(0.015)	(0.131)	(2.344)	(0.085)	(3.641)	(0.035)
C. Full Sample							
(1)	---	---	---	-2.622	---	---	---
				(0.534)			
(2)	---	0.015	0.182	-3.021	-0.008	-1.276	---
		(0.010)	(0.091)	(1.417)	(0.056)	(2.235)	
(3)	---	0.011	161.000	-3.438	-0.018	-1.522	0.052
		(0.010)	(0.091)	(1.405)	(0.055)	(2.192)	(0.023)

Notes. There are 776 spells of welfare (12,663 binary observations in the estimation sample) in sample A. There are 615 welfare spells in the sample (10,694 binary outcomes) in sample B. There are 1,107 spells (18,426 binary outcomes) in sample C.

TABLE 5. SIMULATION OF THE EFFECTS OF PARENTAL AND ETHNIC WELFARE PROPENSITIES ON THE DURATION OF WELFARE SPELLS AND ON EXPECTED PROPORTION OF DURATIONS CENSORED AT 52 MONTHS

		Model 1		Model 2		Model 3	
Sample/Variable	Value	$E(T^*)$	% > 52	$E(T^*)$	% > 52	$E(T^*)$	% > 52
A. Lived with Parents							
Parental Welfare	0.000	17.346	9.6	17.416	9.6	17.408	9.6
	1.000	22.192	17.7	22.029	17.4	21.854	17.0
Ethnic Welfare	0.144	19.095	12.3	19.082	12.2	19.015	12.1
	0.244	22.229	17.8	25.323	23.9	25.731	24.7
B. Female, Lived with Parents							
Parental Welfare	0.000	19.215	12.3	19.151	12.2	19.154	12.2
	1.000	24.162	21.4	23.908	20.8	23.895	20.8
Ethnic Welfare	0.149	21.041	15.4	20.905	15.1	20.903	15.1
	0.249	25.465	24.0	28.895	31.5	28.995	31.8
C. Full Sample							
Ethnic Welfare	0.140	19.337	13.9	19.268	13.6	19.006	13.0
	0.240	23.661	21.9	24.248	22.9	24.664	23.6

Notes: $E(T^*) = \sum_{j=1}^{52} P(T=j) \cdot j + 52 \cdot P(T>52)$, the lower-bound approximation of the mean duration. The remaining variables are evaluated at sample means.

Figure 1. Second-Stage Regressions of Group Fixed Effect on Welfare Use in Parental Generation

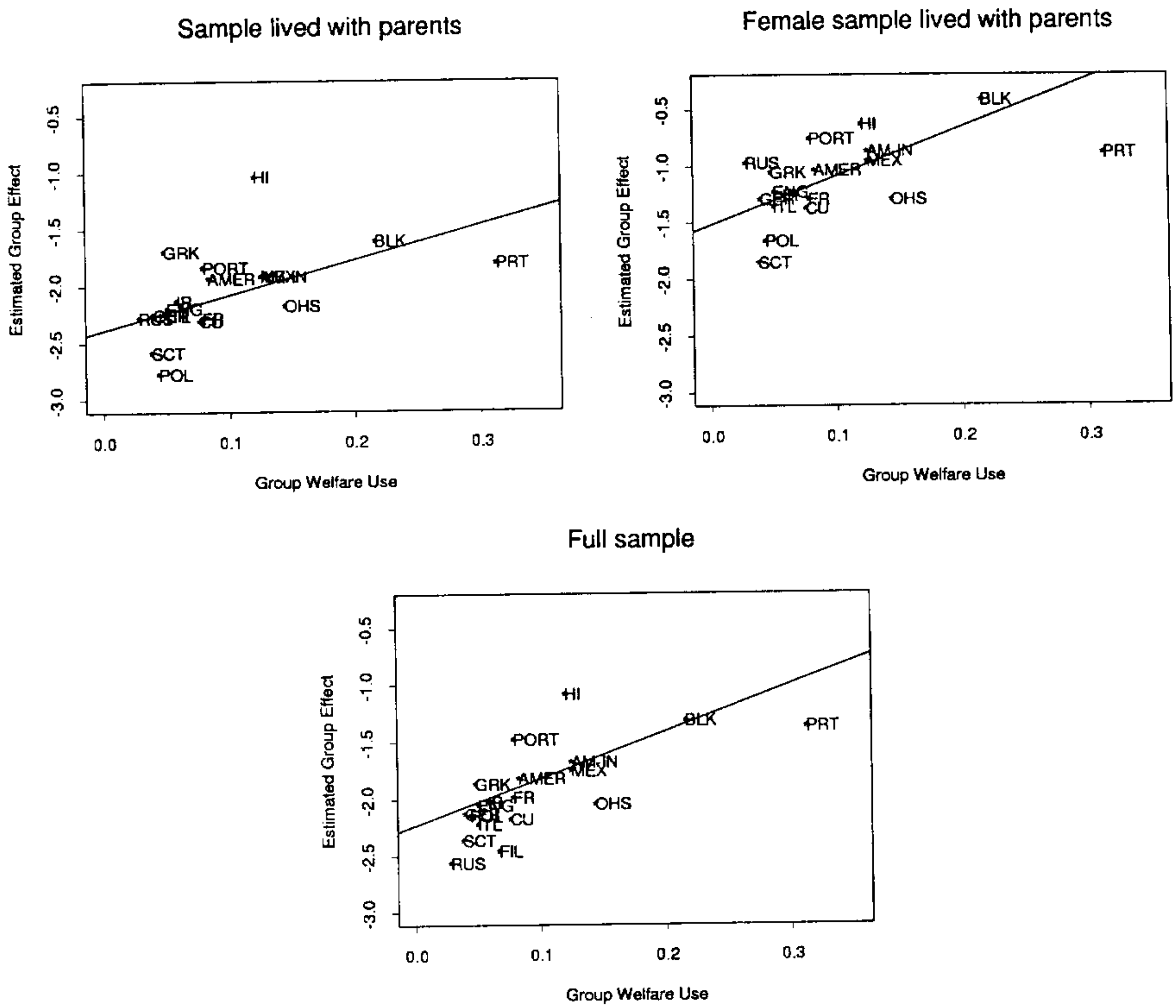

Figure 2. Second-Stage Regressions of Group Hazard Fixed Effect on Welfare Use in Parental Generation