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Eliminating Credit Barriers To Increase Homeownership: How Far Can We Go?

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Abstract

Using data from the 1998 Survey of Consumer Finances (SCF), this paper estimates homeownership rates that would prevail in the United States if borrowing constraints were eliminated. These estimates are obtained by first identifying a group of unconstrained households in the SCF based on a unique set of survey questions. Housing tenure preferences are then estimated over unconstrained households controlling for sample selection effects, and results are used to forecast owner-occupancy rates that would prevail in the absence of borrowing constraints. If all borrowing constraints were removed, *ceteris paribus,* the owner-occupancy rate among non-farm families in the United States would increase just over four percentage points. In contrast, borrowing constraints have a comparatively small effect on the percentage of current renters who expect to own a home in the next 10 years. This pattern suggests that borrowing constraints depress owner-occupancy rates primarily by delaying homeownership rather than permanently excluding families from owner-occupied housing. Additional analysis confirms that household social and financial stability, ability to care for a property, and wealth are all very important determinants of whether families prefer to own. On balance, therefore, evidence suggests that there is room for further relaxation of borrowing constraints to expand access to homeownership. However, such efforts are likely to be more fruitful if attention is given to mortgage products that alleviate borrowing constraints in the early years of the mortgage and if mortgage product innovation is coupled with policies designed to enhance the social and financial stability of families.

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I. Introduction

Few symbols of personal economic success loom larger in the minds of Americans than owning one's own home. Norms favoring homeownership have been further buttressed by the belief that because homeownership is a site-specific investment homeowners take better care of their neighborhoods and, therefore, make good citizens (e.g., Rohe, McCarthy, and Van Zandt 2000; DiPasquale and Glaeser 1999).[1] Although evidence on whether homeowners make better neighbors is still tentative, it is certain that as a society we value homeownership for both personal and social reasons.[2] This is clear from a long history of tax policies that encourage homeownership by reducing the cost of housing for owner-occupiers (e.g., Rosen 1979, 1985). In addition, the last decade has witnessed a series of government and industry efforts designed to reduce mortgage borrowing constraints for low-income families and other disadvantaged groups. These efforts have focused primarily on relaxing wealth requirements and have culminated in the recent introduction of zero and near-zero down payment mortgages for eligible borrowers.[3]

Innovations in affordable mortgage lending—along with the dramatic economic expansion of the 1990s—have helped to raise U.S. homeownership rates to historic levels, from 64 percent in 1989 to just over 67 percent by 2000. Against that backdrop, an important question for both government and the financial industry is as follows: To what extent can further relaxation of borrowing constraints boost homeownership rates, both for the U.S. population overall and for families of different age, race, ethnicity, and financial status? This paper seeks to answer that question by estimating homeownership rates that would prevail

[1]A growing literature on social capital has examined whether homeowners invest in social capital that would serve to enhance the vitality of their neighborhoods (e.g., Rohe et al. 2000; DiPasquale and Glaeser 1999). For example, do homeowners vote more often than renters? Do they donate more of their time to community functions? Do they take better care of their homes? All such behavior imparts positive benefits on the neighborhood and society in general.

[2]It is difficult to show that homeowners make good neighbors because of a strong simultaneity problem. Families that have a taste for investing in their local neighborhoods likely expect to remain in their community for an extended time. Yet the cost of owner-occupied housing declines with length of stay because the high costs of moving to and from owner-occupied housing can be spread out over a longer period (e.g., Rosenthal 1988). For that reason, families intending to remain stationary for a long time are more likely to own their home, *ceteris paribus*, and, for the same reason, are more likely to invest in a variety of forms of neighborhood social capital.

[3]For example, Zero Down™ is an affordable mortgage product offered by Bank of America; it is available in 23 states and Washington, DC. It is a conventional mortgage that requires zero down payment. In addition, closing costs can come from a gift or the seller, or can be financed (see Bank of America 1998). These loans and other affordable mortgage products are typically issued only to individuals with low or moderate income relative to the areas in which they live. Moreover, as a broad characterization, the more relaxed the underwriting standards, the more strict the eligibility criteria with regard to credit risk that govern whether a prospective borrower would have access to the mortgage product.

for various subgroups of the population if all borrowing constraints were eliminated, but households otherwise abide by their budget constraints.

Addressing these issues is difficult because of three fundamental problems that all studies on the impact of borrowing constraints face. Researchers must identify which families are credit constrained; they must evaluate how those families would behave if borrowing constraints were relaxed, *ceteris paribus*; and they must control for the myriad possible constraints that lenders impose on prospective borrowers. In the context of homeownership, such constraints include down payment standards, as well as house payment–to-income and total debt payment–to-income ratios, and the various ways these standards are applied for different types of loans (e.g., fixed versus variable-rate mortgages). In addition, to the extent that nonmortgage debt like auto and consumer loans can be used as substitutes for mortgage debt, borrowing constraints outside of the mortgage market potentially affect housing tenure decisions as well.

The solution to these problems offered here is to apply sample selection methods to unusually rich data from the 1998 Survey of Consumer Finances (SCF).[4] At the core of the estimation strategy is a unique set of survey questions that enable one to identify *a priori* a group of households who almost certainly hold as much mortgage debt as they would like given prevailing market rates and the menu of existing mortgage products. These families are characterized as not credit constrained (unconstrained) in the analysis to come, while all other households are characterized as *possibly* credit constrained for reasons that will become apparent later in the paper. Housing tenure preferences—to own or to rent—are then estimated over just the unconstrained families, controlling for the endogenous selection of families into the unconstrained group through a three-celled bivariate probit model. The resulting housing tenure coefficients reflect the impact of household demographic and financial characteristics on preferences for owning when families are subject to their budget constraints but not subject to binding borrowing constraints. Those coefficients are used to predict homeownership rates that would prevail for the entire population—unconstrained and possibly credit-constrained families. Comparing predicted with actual homeownership rates permits one to evaluate how much higher homeownership rates would likely rise if all borrowing constraints were eliminated, *ceteris paribus.*

Is there opportunity for industry and government to expand homeownership through further relaxation of borrowing constraints? Results from the analysis suggest a qualified yes.

[4]The principal limitation of the SCF is that strict regulations designed to protect confidentiality do not allow the researcher to observe any information related to the geographic location of the household. However, as will become apparent, the density of development in the household's neighborhood is available, and this variable acts as an excellent proxy for central city status. Additional details on the SCF data are provided later in the paper.

If all borrowing constraints were suddenly removed, *ceteris paribus,* and all households could instantly change their housing tenure if they so chose, the owner-occupancy rate among non-farm families in the United States would increase by just over four percentage points. Not surprisingly, these impacts are distributed unequally across different subgroups of the population. Borrowing constraints have far more impact on homeownership rates among low-income families than any other group, roughly 10 percentage points. Opportunities to expand homeownership also exist among young and middle-aged families for whom elimination of borrowing constraints would boost homeownership rates by roughly seven percentage points. In addition, borrowing constraints appear to have more impact on Hispanic homeownership rates relative to whites, while the evidence is mixed with regard to effects on African Americans.

To further explore these findings, two modifications are made to the model to shed light on a different but related question: To what extent do borrowing constraints serve to delay—as opposed to permanently exclude—access to owner-occupied housing? To examine this question, current housing tenure status is replaced with a survey question that inquires whether families *expect* to own a home in the next five to 10 years. In addition, the model is estimated only over current renters. If borrowing constraints have no effect on renter expectations of future owner-occupancy status, then those families that currently rent because of borrowing constraints must expect to overcome those constraints within a decade.

Results indicate that the percentage of renters who expect to own in the next decade rises by 7.56 percentage points with the elimination of borrowing constraints. Assuming renter expectations are fully realized, that figure translates into 2.47 percentage point increase in the share of current households (owners and renters combined) that eventually attain homeownership, since renters account for one-third of all families. In contrast, elimination of borrowing constraints increases owner-occupancy rates by four percentage points as discussed above. The comparatively small effect on renter expectations suggests that borrowing constraints depress current owner-occupancy rates at least in part by delaying access to owner-occupied housing, consistent with recent findings by Goodman and Nichols (1997). Government and industry efforts to expand homeownership, therefore, may want to focus on mortgage product designs that alleviate borrowing constraints primarily in the early years of a mortgage to encourage earlier access to owner-occupied housing.[5]

The remainder of the paper is organized as follows. The next section reviews selected portions of the literature that bear on discussions of the impact of borrowing constraints on

[5]For example, graduate payment mortgages (GPMs) and price level adjusted mortgages (PLAMs) reduce tilt problems by allowing for increasing nominal payments over the life of the mortgage.

homeownership. That literature is placed in the context of recent policy debates that have affected the operation of the mortgage market during the 1990s. As will become apparent, those debates also influenced the direction of academic research in this area. Section III describes the empirical model. Section IV presents the data, and Sections V and VI present the results from the bivariate probit analyses and simulations. Section VII concludes.

II. Literature Review and Policy Context: The Impact of Borrowing Constraints on Homeownership

Why Lenders Ration Credit with Down Payments and Other Constraints

Before proceeding further, it is useful to clarify why lenders would choose to ration credit through down payments and other underwriting criteria as opposed to simply using the loan rate to clear the market. Imperfect competition coupled with racial discrimination is certainly one motivation for such behavior. More generally though, Stiglitz and Weiss (1981) provide much of the theoretical foundation for why competitive lenders would choose to ration credit through terms other than the loan rate. They argue that moral hazard, adverse selection, and asymmetric information between borrowers and lenders regarding credit risk can give rise to equilibrium credit rationing in which loan rates may be set at below-market clearing levels. Such an outcome arises because information is costly and lenders have an imperfect ability to classify borrowers according to default risk. Under such conditions, lenders price loans based on the expected return on the loan portfolio rather than the expected return on the individual loans. The expected return on the pool of loans, however, depends both on interest earnings on loan payments and on expected default costs, each of which rises with the loan rate. In the latter case, as the loan rate increases borrowers have an incentive to invest in riskier projects (moral hazard) with higher expected returns. In addition, as loan rates rise, prospective borrowers with strong aversions to default tend to drop out of the applicant pool first (adverse selection), raising the average propensity to default of the remaining pool of borrowers. Competitive lenders respond to such effects by setting interest rates lower than would occur in the absence of moral hazard and adverse selection effects. Under such conditions, Stiglitz and Weiss (1981) show that it is possible that the competitive equilibrium will occur at below-market clearing interest rates.[6]

In an extension of the initial model, Stiglitz and Weiss (1981, Part IV) further describe a red-lining model in which lenders vary loan rates across borrowers on the basis of observable differences in credit risk—a model that fits more closely mortgage markets in practice. That model still allows for an equilibrium in which sufficiently high-risk loan

[6]Duca and Rosenthal (1991) provide empirical support for these ideas.

applicants may be credit constrained because of adverse selection and moral hazard. Duca and Rosenthal (1994b) extend that model further and argue that regulatory restraints such as fair lending laws may have had the unintended effect of increasing the degree to which lenders use nonrate terms to control for perceived differences in applicant credit risk. This could arise if lenders choose to offer the same mortgage rates to borrowers of different risk attributes to reduce the likelihood of costly discrimination suits.[7] Under such conditions, lenders would have an incentive to control for perceived differences in credit risk through less visible means, such as nonprice constraints like down payment standards.

One implication of these arguments is that as information becomes increasingly plentiful and inexpensive to obtain, lenders will be increasingly able to adjust individual loans—either through interest rate or nonrate terms—to meet the characteristics of individual applicants. This suggests that with the information technology revolution, we should expect the degree to which lenders rely on nonrate terms to diminish, which is what appears to have occurred in the mortgage market throughout the 1990s.

The Early Empirical Literature on Credit Rationing

The earliest empirical work that has direct bearing on the design of the present paper comes from research testing the robustness of the life-cycle permanent income hypothesis (LCPIH). That literature includes a number of important cross-sectional and panel data analyses, such as those by Hall and Mishkin (1982), Hayashi (1985), and Zeldes (1989). These papers provide evidence that the time path of consumption expenditures for households that are *not* credit constrained differs from that of families for whom borrowing constraints *may* be binding. Based on these findings, authors argued that borrowing constraints were binding for many people, a violation of the LCPIH. A limitation of these studies, however, is that the data used do not directly identify credit-constrained and unconstrained families. Instead, the studies above assume that families with either high wealth-to-income ratios or high savings rates are not credit constrained. Given that the demand for debt increases with wealth and income in response to increased demand for consumer durables such as housing, high-income and high-wealth families could still be credit constrained.[8] This concern has raised questions about whether the analyses above suffered from coding errors when splitting the sample on the basis of who is not credit constrained.

In response, several studies in the early 1990s drew on the Federal Reserve's Survey of Consumer Finances (SCF). That survey allows the researcher to directly identify families

[7]Using 1983 Survey of Consumer Finances (SCF) data, Duca and Rosenthal (1994b) found no evidence that lenders vary loan rates across borrowers on the basis of observable differences in risk attributes.

[8]8Jappelli (1990) and Duca and Rosenthal (1993) provide evidence on this point.

that have recently been turned down for credit, received smaller-than-desired loans, or have been dissuaded from applying for credit. Using the SCF, Jappelli (1990) investigated the characteristics of credit-constrained families, and Cox and Jappelli (1993) estimated the extent to which borrowing constraints reduced the levels of debt held by such families.[9] These two studies, however, did not control for a number of variables used by lenders in evaluating loan applications, including credit history in the case of Jappelli and both credit history and wealth in the case of Cox and Jappelli.

The importance of controlling for wealth and credit history when analyzing household access to credit has been underscored by a vigorous debate over the last decade about whether racial discrimination restricts the ability of minority households to obtain credit. Although that controversy dates back at least to the 1970s when a wave of fair lending legislation was enacted, the debate became especially sharp following the publication of several sets of studies in the late 1980s. The first of these were newspaper articles that focused attention on the fact that minority populations had much more limited access to mortgage credit than majority white loan applicants, in part because of discrimination.[10] Numerous press reports shortly thereafter focused attention on data from the Home Mortgage Disclosure Act (HMDA).[11] Those data showed that mortgage application rejection rates for African Americans in 1990 were 2.4 times as large as those for white families with similar income (see Canner and Smith [1991] for a detailed description of the HMDA data).

As noted by Rehm (1991b), evidence from HMDA prompted House Banking Committee Chairman Henry Gonzalez to ask "top regulators for an 'immediate' report on what their agencies plan to do 'to correct the lending problems revealed [by the HMDA data].'" However, the HMDA data do not include household credit history or wealth, in addition to other important variables that appear on loan application forms. As a result, many other individuals in government, the banking industry, and academia questioned whether the HMDA data implied that lenders discriminate against minority loan applicants.[12] Partly in response to that debate, the Federal Reserve Bank of Boston conducted a landmark study of

[9] See also papers by Duca and Rosenthal (1993, 1994a).

[10] In May 1988, the *Atlanta Constitution* published a four part series, "The Color of Money," while the *Detroit Free Press* published a similar series in July 1988.

[11] Beginning in 1990, lenders were required by HMDA to report the location of residential loans made along with the income, race, and gender of loan applicants and whether the loan application was withdrawn (by the applicant), approved, or denied. See Rehm (1991a, 1991b) and Munnell et al. (1996) for further discussion of the HMDA data.

[12] For example, Rehm (1991b) notes that although Governor John LaWare of the Federal Reserve described the HMDA data as "very worrisome," he indicated that more information was needed. Similarly, Rehm (1991b) reports that "Leading industry groups, such as the American Bankers Association, have maintained that the Fed data do not take into account information crucial to credit decisions, such as a loan applicant's credit history, other debts, or employment."

mortgage application denial rates in Boston (Munnell et al. 1996) using a much wider range of loan applicant characteristics than previously analyzed. An important finding of the study was that allowing for differences in loan applicant wealth and credit history reduced but did not eliminate race-related differences in mortgage denial rates. Subsequent exhaustive analyses have largely supported these claims (e.g., Turner et al. 1999; Turner and Skidmore 1999).[13]

Who Chooses to Own Versus Who Has the Ability to Own?

Recently, the literature on the impact of borrowing constraints on homeownership appears to have split along two lines that address related but different questions: who would *choose* to own a home under different underwriting criteria, and who would have the *ability* to purchase a home under different underwriting criteria? Studies focusing on the former question have their beginnings with papers by Linneman and Wachter (1989) and Zorn (1989). These studies identified credit-constrained households by assuming that prospective homeowners with more than a 28 percent house payments–to-income ratio were credit constrained.[14] However, many lenders in the 1980s had house payments–to-income limits of more than 28 percent, while families with a bad credit history were likely to face tighter-than-average credit standards (e.g., Boyes, Hoffman, and Lowe 1989; Trans Data Corporation 1986).[15] Nevertheless, an important finding from Linneman and Wachter (1989) is that down payment constraints appear to restrict access to homeownership with greater frequency than income does. More recently, papers by Quercia, McCarthy, and Wachter (1998) and Haurin, Hendershott, and Wachter (1997) have become quite sophisticated in the manner in which the sample is stratified into constrained and unconstrained households, and the degree to which the models examine the impact of a wide range of different possible underwriting criteria. An important finding from these studies is that borrowing constraints

[13]In response to these findings and related community pressure, the Federal Reserve Board approved several large bank mergers conditional on the requirement that merger applicants meet lending goals in minority neighborhoods (see for example, the description of Bank of America's merger with Security Pacific in Thomas [1992], p. A6). In addition, Fannie Mae and Freddie Mac, the dominant players in the secondary mortgage market, established new low-down-payment loan programs designed to reduce credit barriers for underserved populations (e.g., Reuters 1991). Private industry responded to these programs by originating new mortgage products targeted at previously underserved groups knowing that they could then sell such products to the secondary market. See Listokin and Wyly (2000) for a careful review of affordable mortgage products currently available on the market and the history of how those products came to be.

[14]This assumption derives, presumably, from secondary mortgage market criteria at the time that generally prohibited the securitization of mortgages with house payments–to-income ratios of more than 28 percent.

[15]Trans Data Corporation data on official credit standards for primary mortgage lenders across the United States (in 1986) indicated that many lenders had house payments–to-income limits above 28 percent.

continue to impede homeownership for underserved groups in the population, including younger families, minorities, and low-income households.[16]

In contrast to these papers, housing "affordability" studies by Savage (1999), Listokin et al. (1999), and others have focused more on the question of who has the ability to own a home.[17] As a broad characterization, these studies proceed by first specifying a reference value home for each family in the sample. The reference home is most often specified as a function of the distribution of owner-occupied house values in the household's region of the country (e.g., the 10th or 25th percentile).[18] Next, data are examined to determine whether or not individual families have sufficient income and wealth to satisfy underwriting guidelines for a range of different prospective mortgage products. Based on that exercise, the percentage of renters capable of buying the reference home under different underwriting criteria is determined.[19]

The housing affordability approach has made an important contribution, and the Census Bureau affordability tables are a valuable resource for housing analysts. That said, in the context of the present paper it is important to bear in mind that mortgage market constraints are but one of a number of important factors governing whether families would choose to own a home. Perhaps the most general framework in this regard is the theoretical model of Henderson and Ioannides (1983), and subsequent empirical support for the model in Ioannides and Rosenthal (1994). In that model, families have both a consumption and investment demand for housing. Consumption demand is sensitive to the demand for shelter. Investment demand is driven by portfolio considerations. If investment demand exceeds consumption demand, the family could choose to own a home equal to investment demand and rent out the unwanted space: in this case the family is better off if it owns. Alternatively, if consumption demand exceeds investment demand, the family would not want to purchase housing up to the level of consumption demand since that would constitute a bad investment: in this case the family is better off if it satisfies its consumption demand by choosing to rent its principal residence.

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[16]Mayer and Engelhardt (1996) provide additional evidence that many new homebuyers receive gifts of one sort or another in the few years prior to purchase of the first home, especially among families likely to be credit constrained. See also Haurin, Hendershott, and Wachter (1996) for a related discussion.

[17]The housing affordability approach is also explicit in the National Association of Home Builders (NAHB) housing opportunity index (HOI) and the National Association of Realtors (NAR) housing affordability index (HAI).

[18]Listokin et al. (1999) also estimate the preferred house value for individual households and then compare the application of that reference house to reference values based on a percentile of the house value distribution as described above.

[19]Studies by Savage (1999), Listokin et al. (1999), and affordability tables at the Census Bureau web site (http://www.census.gov/hhes/www/hsgaffrd.html) have all been based on data from the Survey of Income Program Participation (SIPP).

The Henderson-Ioannides model, while stylized, offers considerable guidance on how we might want to think about the question of who prefers to own a home in the absence of binding borrowing constraints. It is well documented that moving from owner-occupied housing is far more expensive than moving from rental housing because of Realtor fees, legal fees, and taxes (e.g., Rosenthal 1988). For that reason, families who do not expect to move soon are more likely to prefer to own because they can spread the high transactions costs of moving to and from owner-occupied housing over a longer time. Consider, however, that family and financial instability both increase the frequency with which a family moves, reducing the return on owner-occupied housing. In the context of the Henderson-Ioannides model, such instability lowers the investment demand for housing. For these families, owning a home could be a bad investment, and many such families may prefer to rent.

As also argued by Henderson and Ioannides (1983), there is a tendency for individuals who have a taste for maintaining their dwellings to be undercompensated for such behavior in the rental market. Thus, for families that are good at maintaining their homes, owner-occupied housing may be a better investment, causing investment demand to be high; these families may prefer to own. Finally, Fu (1991) modifies the Henderson-Ioannides model to evaluate how household wealth affects preferences for owning. Suppose that housing consumption is a normal good and absolute risk aversion declines with wealth. Then, since owner-occupied housing is a risky investment, housing investment demand likely increases faster with wealth than housing consumption demand. As a result, wealthy families are more likely to prefer owner-occupied housing even in a setting free of taxes and financing issues.[20]

Given that the goal of this paper is to estimate the homeownership rate that would prevail if borrowing constraints were eliminated, it is essential to take into account all the determinants of whether a family would choose to own a home. The discussion above suggests that household social and financial stability—important determinants of household mobility—are factors that must be considered. In addition, the family's ability to maintain its home and the family's level of wealth—an important determinant of the level of risk the family is willing to accept—are important as well. Finally, to the extent that owning a home has become a societal norm, this suggests that households value owning a home for reasons unrelated to financial gain (in contrast to the Henderson-Ioannides [1983] model). Thus, anything that might enhance the degree to which owning a home directly affects a family's sense of well-being also belongs in models of housing-tenure preferences. These principles will guide the variable selection in the empirical work to follow.

[20] Ioannides and Rosenthal (1994) provide empirical support for the idea that housing investment demand is indeed more sensitive to wealth than consumption demand.

III. Empirical Model

Overview

In the most general setting, families choose their housing tenure to maximize utility subject to two sets of constraints: (1) their intertemporal budget constraint and (2) borrowing constraints imposed by lenders. Although the budget constraint is binding for all families provided that savings and bequests are treated as future consumption, borrowing constraints are binding only for a subset of households as the previous section outlined. The focus of this paper, of course, is to determine housing tenure preferences and homeownership rates subject only to the household's budget constraint. To pursue that goal, the model below is developed in a manner that is tailored to the characteristics of the Survey of Consumer of Finances (SCF), three features of which are important to emphasize here.

First, the SCF is a cross section of individual households. In that regard, all families face the same macroeconomic conditions, though access to different mortgage products varies with household demographic and financial characteristics. Second, the SCF permits one to identify *a priori* a group of households for whom borrowing constraints are not binding, taking into account all possible credit constraints within and outside of the mortgage market. These families are referred to as *unconstrained* for the remainder of this paper. All other families may or may not face binding borrowing constraints for reasons that will be clarified in Section IV. These families are referred to as *possibly constrained* for the remainder of this paper. Third, survey questions in the SCF permit one to examine two different variables that shed light on housing tenure preferences, whether families currently own or rent, and whether current renters expect to own in the next five to 1- years. When analyzing current housing tenure, all households are included in the sample. When analyzing whether renters expect to own, only renters are included in the sample. Apart from those differences, as will become apparent, the structure of the analysis is identical in both cases. For that reason, the model below is developed only in the context of current housing tenure status, but the reader should keep in mind that this model is also used to estimate whether current renters expect to own as well.

Housing Tenure Preferences When Borrowing Constraints Are Not Binding

Define an unobservable index $\mathrm{I}_{\mathrm{own}}$ that represents the difference in utility between owning and renting when families are subject only to their budget constraints,

$$\mathrm{I}_{\mathrm{own}} = xa_x + Ma_m + \mathrm{e}_{\mathrm{own}}\,. \tag{3.1}$$

This equation determines a family's preferred tenure status *in the absence of binding borrowing constraints*. Elements of x include all demographic and financial characteristics of the household that influence tenure preferences as described in Section II. In addition, elements of M include characteristics of the preferred mortgage product that the household would choose if it were to own, such as fixed versus variable rate, loan rate, amortization period, down payment ratio, etc.

What determines m? Macroeconomic conditions that are identical for all households given the cross-sectional nature of the data, and household characteristics that influence a family's preferred form of financing. Accordingly, M is a choice variable that depends on x and can be expressed as

$$M = h(x) ,$$

where the role of macroeconomic conditions is suppressed to simplify notation.[21] Substituting into (3.1),

$$I_{own} = xt + e_{own} . \tag{3.2}$$

Equation (3.2) is a reduced form expression that captures both the direct and indirect effect of household characteristics on housing tenure preferences in the absence of borrowing constraints. However, because I_{own} is unobserved, the analysis below focuses on the observable discrete housing tenure decisions (OWN) corresponding to (3.2),

$$I_{own} > 0 \Rightarrow \text{OWN} = 1 \text{ , } \textit{own home} \tag{3.3}$$

$$I_{own} < 0 \Rightarrow \text{OWN} = 0 \text{ , } \textit{rent home}$$

where OWN equals 1 if the family owns and 0 if the family rents.

Denote now a second unobservable index that governs whether a family belongs to the unconstrained group or the possibly constrained group, I_{NotCC},

$$I_{NotCC} = xc + e_{NotCC} . \tag{3.4}$$

The discrete observable realizations corresponding to (3.4) are given by

$$I_{NotCC} > 0 \Rightarrow \text{NotCC} = 1, \textit{ unconstrained} \tag{3.5}$$

$$I_{NotCC} < 0 \Rightarrow \text{NotCC} = 0 \text{ , } \textit{possibly constrained}$$

where NotCC equals 1 if the family belongs to the unconstrained group and 0 if the family belongs to the possibly constrained group.

In viewing expressions (3.1–3.5), it is important to recognize that whereas NotCC is observed regardless of whether OWN takes on a value of 1 or 0, housing tenure preferences *free of borrowing constraints* are observed only for families belonging to the unconstrained group: NotCC equal to 1. As is well established in the discrete choice literature (e.g., Maddala 1983), if e_{own} and e_{NotCC} are uncorrelated, observing OWN only for NotCC equal to 1 presents few difficulties. Assuming e_{own} follows a unit normal distribution, one could

[21] Macroeconomic conditions common to all households are captured in the model's constant term.

obtain unbiased and consistent estimates of *t*—the housing tenure preferences in (3.2)—by running a univariate probit model over just that portion of the sample for which NotCC = 1.[22] More generally, however, common omitted variables that influence both the likelihood that NotCC equals 1 and the likelihood that OWN equals 1 would cause estimates of *t* to suffer from sample selection bias owing to the endogenous character of the sample selection procedure. To allow for this possibility, a more general estimation procedure is needed.

The Bivariate Probit Model with Three Cells

To avoid sample selection bias, it is necessary to control for correlation between the error terms in the two latent indexes, e_{NotCC} and e_{own}. If (3.2) could be estimated directly, a common approach would be to use well-known Heckman two-step procedures by augmenting (3.2) with a Mills ratio term based on first-stage probit estimates of (3.5). Subject to identification conditions and functional form, including the Mills ratio enables one to obtain consistent estimates of *t*. In the present context, however, I_{own} is not directly observable. Instead, the discrete variable OWN is observed. In this case, a nonlinear analogue of the Heckman procedure is to estimate a bivariate probit model over (3.3) and (3.5) with just three cells, where OWN is observed only for NotCC = 1 as noted above.

More formally, assume that e_{NotCC} and e_{own} follow a bivariate standard normal distribution with mean zero and covariance $\sigma_{NotCC,own}$.[23] Then the log likelihood function (*L*) for this model is given by,

$$L = \sum\{(1\text{-NotCC})\cdot\log[F(-xc)] + \text{NotCC}\cdot\text{OWN}\cdot\log[G(xt,xc,\sigma_{NotCC,own})] \qquad (3.6)$$
$$+ \text{NotCC}\cdot(1\text{-OWN})\cdot\log[G(-xt,xc,-\sigma_{NotCC,own})]\},$$

where *F(·)* and *G(·)* are the standard unit and bivariate normal distributions, respectively, and the log-likelihood function is evaluated separately for all observations in the entire sample, *i* = 1, …, I.[24] Note, however, that whereas each observation in the sample contributes to the identification of *c*, the parameters governing whether a family belongs to the unconstrained group, only those families for which NotCC is equal to 1 contribute to identification of *t*, the parameters governing housing tenure preferences. In addition, sample selection effects are controlled for because the covariance between e_{NotCC} and e_{own} appears in the last two

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[22]Note that when e_{NotCC} is independent of e_{own}, the expected value of I_{own} is *xt* since $E[e_{own}|e_{NotCC}] = 0$, and consistent estimates of *t* can be obtained by running a univariate probit model on (3.3) using only unconstrained families.
[23]The variances of e_{NotCC} and e_{Own} are normalized to 1 because the parameters of the bivariate probit model can be estimated only up to a scale factor. See Maddala (1983) for further discussion.
[24]Boyes, Hoffman, and Low (1989) estimate a similar three-celled bivariate probit model for the credit card market.

bracketed terms of (3.6) and is simultaneously estimated along with *t* and *c*. Thus, (3.6) provides unbiased and consistent estimates of *t*.[25]

Simulating Owner-Occupancy Rates in the Absence of Borrowing Constraints

Estimates of *t* obtained from (3.6) reflect the impact of household financial and demographic characteristics on housing tenure preferences in the absence of borrowing constraints. Given such estimates, it is possible to simulate the percentage of the population that would choose to own by computing the mean of $F(xt)$ over the entire sample, where $F(\cdot)$ is the unit normal distribution function as previously noted. Comparing that estimate to the actual frequency of owner-occupiers gives an estimate of the impact of borrowing constraints on homeownership rates. Repeating the simulation exercise for various subsets of the population permits evaluation of the impact of borrowing constraints on different subgroups by race, income, age, etc.

IV. Data

The data used to estimate the model are taken from the 1998 Survey of Consumer Finances (SCF), a household-level survey designed to provide unusually rich information on the financial characteristics of households. In total the SCF provides data on roughly 4,300 households.[26] Of these, roughly 2,800 are selected so as to be representative of the entire United States, while the remaining households over-represent wealthy families and are drawn from tax files. To protect confidentiality, the public use version of the 1998 SCF does not allow the analyst to separately identify the representative and tax-based samples. However, sampling weights provided with the data permit one to weight the data such that results are representative of the entire United States. In the work to follow, the bivariate probit model was estimated using unweighted data to obtain the parameter estimates on the assumption that all the covariates in the model are exogenous. The simulations, in contrast, were

[25]An issue of identification does remain. Selection models such as the one above provide more reliable results when there are variables included in the selection equation (equation [3.5] in this case) that do not belong in the equation of interest (equation [3.3]). In the work to follow, as will become apparent, there appear to be several natural exclusion restrictions, most notably a set of credit history variables that clearly belong in the credit model but that have zero coefficients in the tenure preference model. For a more detailed discussion of bivariate probit models with censoring see Maddala (1983) or Tunali (1986).

[26]The SCF data are imputed five times to control for missing values and also to protect the confidentiality of some respondents with especially unusual and highly visible characteristics (such as very high wealth). When estimating the bivariate probit models, all five implicates totaling over 21,000 records were used and the standard errors were divided by the square root of 5 to adjust for the "true" sample size. See the 1998 SCF manual and Kennickell (1998) for details.

calculated using the sampling weights to ensure that the simulation results are representative of the United States.[27]

A special feature of the SCF is that households were asked if they had a request for credit turned down by a particular lender or creditor in the past five years, or had been unable to get as much credit as they had applied for. Households were also asked if there had been any time in the past five years that a person (or the person's spouse) had *thought* about applying for credit at a particular place, but changed his or her mind because the household thought it might be turned down. Based on these questions, a household was classified as possibly credit constrained (NotCC = 0) if at least one of the following three conditions did not hold: (1) the household had not had a loan request turned down, (2) the household had not had a loan request only partially granted, and (3) the household had not initially considered applying for credit but then chose not to because it thought that it would be turned down. If instead *all three of the conditions above* held, then the family was classified as not credit constrained (NotCC = 1).[28]

A further strength of the 1998 SCF is the rich information included on the determinants of household wealth, credit history, and expectations. Thus, in addition to the usual battery of demographic characteristics, a number of household attributes not typically found in most major surveys are included in the model, such as expected income growth, inheritances, gifts and settlements, credit history attributes, and indicators of employment stability. The main limitation of the SCF is that it does not provide information on household location because of strict rules governing confidentiality. However, the SCF provides information on the density of development in the household's neighborhood, and that information serves as an excellent proxy for central city/suburban status.[29] To facilitate review, a description of the principal variables in the model is provided in the tables (see appendix).

A final point concerns sample composition and use of the SCF relative to other data sets. The most widely cited homeownership rates are those from the U.S. Census Bureau. Those estimates are based on the Consumer Population Survey (CPS) for the entire United

[27]See Kennickell (1999) for a careful discussion of the SCF sampling weights.

[28]This is a more demanding definition of who is not credit constrained than was used by Duca and Rosenthal (1994a). In that paper, families that successfully reapplied upon having a loan application rejected or only partially accepted were considered not credit constrained. Classifying such families as NotCC = 0 reduces the efficiency of the estimated housing tenure preferences but increases the likelihood of obtaining unbiased and consistent estimates. The reason is that the three-celled probit model outlined in the previous section requires that one identify a subset of the sample that is clearly not credit constrained. In contrast, the model does not require that everyone in the alternate category be credit constrained. Instead such families *may* be credit constrained, analogous to studies testing the LCPIH by Zeldes (1989) and others.

[29]The SCF reports whether nearby buildings are less than 21 feet apart, 21 to 100 feet, or more than 100 feet apart.

States including both the farm and nonfarm sectors. However, the housing needs and opportunities of individuals living on farms—both farm owners and employees—are arguably rather different than for the nonfarm sector. In addition, much of the policy focus with regard to the creation and marketing of affordable mortgage products has centered on urban areas. For these reasons, the analysis in this paper is based just on the nonfarm portion of the 1998 SCF in order to provide a sharper picture of the nonfarm sector. Estimates in this study, therefore, could potentially differ from published Census Bureau reports for two reasons: use of the SCF data and focus on nonfarm populations. Before proceeding, it is important to clarify the possible effect of these differences.

Figure 1a compares the racial distribution of the population using the CPS and SCF based on samples that are representative of the entire U.S. farm and nonfarm population for the years 1989, 1992, 1995, and 1998. Figure 1b makes a similar comparison of homeownership rates by race.[30] In addition, the last column in both figures reports values for the 1998 SCF based only on the nonfarm population. Bear in mind that all the SCF data are weighted to ensure they represent their respective populations as discussed earlier.

As is apparent in the figures, differences in the reported values between the CPS and SCF data are small and likely reflect differences in the manner in which certain questions are asked regarding homeownership status and race (see Kennickell [1999] for a discussion of this point). This indicates that the sampling weights for the SCF do an excellent job of matching the CPS and that weighted data from the SCF are representative of the United States. A more substantiative difference arises when comparing the combined farm plus nonfarm populations to just that of the nonfarm population for 1998 using just the SCF. On the one hand, the racial distribution of the population is little different in the last two columns of Figure 1a. On the other hand, for each subset of the population other than Hispanic (for which there is little difference), the homeownership rate in Figure 1b is more than one percentage point higher for the nonfarm population than for the combined farm plus nonfarm population. Those differences boost the overall homeownership rate from 66.1 percent in 1998 for the farm plus nonfarm sector to 67.4 percent for the nonfarm sector. It is important to bear in mind, therefore, that the base homeowership rate for the study group in this paper is 1.3 percentage points higher than is commonly cited in the U.S. Census reports.

[30]The combined farm plus nonfarm values are taken from Kennickell (1999).

V. Results

Summary Statistics

Summary statistics of all the variables included in the bivariate probit regressions are provided in Figure 2 for the full sample and for various subsets of the population. As before, all values are weighted to ensure they are representative of the United States in 1998.

In the top row of the figure, observe that homeownership rates vary widely not just across race, but also with income, age, and location. Among families whose total household income is in the first decile, just 34.4 percent own their homes, rising to 60.2 percent for families whose incomes fall within the 25th to 50th percentiles. Only 40.7 percent of families with household heads under age 35 own their homes. Among families living in densely developed areas (areas where the nearby buildings are within 21 feet of each other), only 54.4 percent own their homes.

The second row in Figure 2 reports the percentage of families that currently rent but expect to buy a home in the next five to ten years. Overall, 8.7 percent of U.S. families belong to this category. Not surprisingly, however, such families are disproportionately found among households under age 35; 22.1 percent of this group are renters that expect to own in the next decade. African American and Hispanic families also include a higher share of renters that expect to own in the coming decade, but this arises because of the greater frequency of renters among these populations. As will become apparent shortly (in figure 4), among renters, the frequency of families that anticipate owning in the next ten years is similar for whites, African Americans, and Hispanics.

Interpreting Estimates from the Bivariate Probit Models

Figure 3 presents estimates for the two versions of the bivariate probit model discussed earlier, first for both renters and owners with *current owner-occupancy status* as the housing tenure variable, and then again for just renters with *expect to own in the next five to ten years* as the tenure variable. Because of the nonlinearity of the bivariate probit model, coefficient estimates from the model can be used only to evaluate the sign of the estimated relationship. To facilitate interpretation, therefore, Figure 3 presents estimates of the partial derivatives for the covariates instead of the original model coefficients. Those derivatives are calculated as

$$\theta_{partial} = \theta \cdot [\sum w_i \cdot f(x_i\theta)] / \sum w_i \,,$$

where θ is the vector of probit model coefficients for the tenure and NotCC equations ($\theta = t, c$), $f(x_i\theta)$ is the unit normal density function evaluated at $x_i\theta$, w_i is the sampling weight for observation i, and $\sum w_i$ is the sample size (appropriately weighted). Calculating $\theta_{partial}$ in this manner ensures that the partial effects are representative of the United States while permitting one to interpret the partials as for a linear probability model.[31]

The Credit Constraint Equation

Recall that the credit-constraint equation is included in the model primarily to control for sample selection effects. In that regard, results from the NotCC equation are of secondary importance. Accordingly, discussion of the NotCC results below is brief and focuses on certain key variables that help to ensure that the NotCC equation serves the function for which it is intended.[32]

At the bottom of Figure 3, observe that a past history of bankruptcy and having made loan payments more than two months late in the previous year both have no influence on preferences for current owner-occupancy status (column 3). On the other hand, those credit variables have highly significant, large negative effects on the likelihood that a family belongs to the unconstrained group. These results indicate that lenders impose tighter underwriting standards on loan applicants with a bad credit history, but that credit history has little effect on housing tenure preferences per se. As such, the credit variables serve as strong exclusion restrictions: they belong in the credit model but do not influence tenure preferences. Such exclusion restrictions reduce co linearity problems that would otherwise limit the ability of the model to control for sample selection effects.

Current Owner-Occupancy Status

Focus now on estimates of the current owner-occupancy equation (column 3) and recall that these estimates measure the impact of household attributes on preferences for living in owner-occupied housing in the absence of binding borrowing constraints, *ceteris paribus.* As discussed earlier, the desire to live in owner-occupied housing is influenced by family and financial stability (which affect expected mobility), ability to care for the home, wealth

[31]As an example, the probability that a family wants to own is 3.5 percentage points higher if the household head is male, as seen in the second column and first row of Figure 3.

[32]Interpreting many of the coefficients in the NotCC equation is difficult because households belong to the unconstrained group if they prefer to hold less debt than lenders are willing to allow (e.g., Duca and Rosenthal 1993). As such, the NotCC coefficients reflect the impact of household attributes on the maximum amount of debt lenders are willing to issue *relative* to a family's demand for debt. For example, receipt of an inheritance (gift or settlement), the dollar value of such a receipt, and the dollar value of future such receipts, all have zero effect on the likelihood that a family belongs to the unconstrained group. That result is consistent with Duca and Rosenthal (1993) who found that wealth does not affect the likelihood of being credit constrained based on 1983 SCF data. This does not, however, imply that lenders care little about loan applicant wealth. Rather, it indicates that the willingness of lenders to issue more debt as household wealth increases is roughly offset by an increase in demand for debt.

(which affects risk aversion), and other attributes that contribute to a family's intrinsic taste for homeownership.

Consider first the role of family stability. Because couples invariably enter into marriage with the expectation of maintaining a stable family, one would expect married families to be more likely to own a home. Indeed, of all the demographic factors, marital status is by far the most important determinant of homeownership: married couples are 20.7 percentage points more likely to prefer to own a home than nonmarried families, *ceteris paribus.*[33] Similarly, it is well documented that young households are more mobile. For these families owner-occupied housing could prove expensive relative to renting. Estimates from the model support that argument. Each additional year of age (for the household head) has little effect on the desire to own up until age 35. After that, however, older households are increasingly likely to prefer to own. An analogous argument can be made with respect to family size. Because it is more traumatic and expensive to move larger families, larger families tend to be less mobile and should, therefore, be more likely to prefer owner-occupied housing. Again, estimates support this argument. With each additional person in the household, families are 2.2 percentage points more likely to prefer to own.

Consider next financial stability. Families for which the head works full time have more secure income than those where the head does not work full time. Of the financial variables in the model, this turns out to be the most important determinant of homeownership. Among families with a head working full time, the likelihood that the family prefers to own is 9.96 percentage points higher than if the head was not working full time. Moreover, after controlling for the employment status of the head (and spouse), total household income has no effect on preferences for owning. Similarly, if the household usually knows what its income will be next year, the family is 5.8 percentage points more likely to want to own. Conversely, as the number of full-time jobs the head has previously held increases—after having already controlled for age of the head—the family is less likely to want to own. The interpretation on this result is that frequent job changes reflect financial uncertainty and increase the likelihood of moving, both of which reduce the appeal of homeownership.

Also, observe that an increase in real income in the previous five years has a positive effect on preferences for owning but is only marginally significant, while expectations that real income will increase in the next five years have a *negative* and significant effect on

[33]Divorced families are also more likely to prefer owning. This may reflect that owning one's home is habit-forming. In addition, capital gains tax provisions still in place in 1998 created financial incentives for individuals to remain owner-occupiers once the first home had been bought (e.g., Hoyt and Rosenthal 1990, 1992).

preferences for owning. The former result may signal increased financial security as the family's income status improves. The latter result could reflect option-type effects. Because housing demand increases with income and moving from owner-occupied housing is costly, if income might increase substantially in the next few years it would make sense to wait before buying until the future income is better known.

Owner-occupied housing also requires more care from the occupant than rental housing. For that reason, one might expect that families in bad health would be less likely to want to live in owner-occupied housing. Results here are intriguing. If the head is in bad health, that has little effect on preferences for living in owner-occupied housing. However, if the spouse is in bad health, that reduces the likelihood that the family prefers to live in owner-occupied housing by 15.6 percentage points. More research is required to sort out exactly what this pattern reflects. One possibility, though, is that the spouse is the principal family member to maintain the home. If the spouse is in poor health and the head is occupied elsewhere (at work, for example), then the family may find that rental housing becomes relatively more attractive.

Owner-occupied housing is a risky, site-specific investment. Because risk aversion tends to diminish with wealth, one would expect that wealthy individuals would be more interested in owning a home (e.g., Henderson and Ioannides 1983, Fu 1991, Ioannides and Rosenthal 1994).[34] Several variables serve as proxy for household wealth and yield results largely supportive of that argument. Families with more highly educated heads are more likely to prefer to own: college degree or more raises the likelihood of owning by 4.5 percentage points relative to a high school degree, while less than a high school degree lowers the rate by 8.8 percentage points. Similarly, families that have received at least one inheritance, gift, or settlement since 1980 are 7.1 percentage points more likely to prefer to own, although interestingly, the dollar value of such receipts has no influence on preferences for ownership. Finally, families living in expensive areas must lever up further to purchase a home for any given level of family wealth. Such investment strategies are risky. In keeping with this argument, families living in neighborhoods where nearby buildings are more than 21 feet apart (characteristic of suburban and rural neighborhoods) are roughly 12 percentage

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[34] The SCF contains very detailed information on household assets and debts and permits one to do an excellent job of calculating net wealth. However, owner-occupied housing is an important asset and, as such, influences the family's portfolio and its wealth. Wealth, therefore, is endogenous and cannot be directly included in the model (see Haurin, Hendershott, and Wachter [1997] for a careful discussion of this point). Because the primary goal of this paper is to forecast homeownership rates that would prevail in the absence of borrowing constraints, exogenous determinants of wealth are included directly in the model rather than attempting to use a two-stage least squares type procedure to include wealth directly in the regression. The reduced form approach taken here is more robust relative to the goals of the paper because no restrictions are placed on the exogenous determinants of wealth.

points more likely to prefer owning than if they were living in areas in which nearby buildings were within 21 feet of each other (characteristic of central city environments).[35]

A last set of variables require some attention. Recall that white homeownership rates are roughly 25 percentage points higher than for African-American and Hispanic families. The model variables discussed above account for most of those differences, reducing the race effects to 8.03 and 8.83 percentage points for the African-American and Hispanic categories, respectively. Nevertheless, one has to question why even these sizable race effects remain. In the case of Hispanics, one possibility is that many of the Hispanic families are recent immigrants, a variable not included in the model (see Coulson [1999], for example). More generally, the estimated race effects likely reflect the influence of omitted determinants of financial and family stability, ability to care for the home, wealth, and discriminatory treatment, all of which are correlated with race and ethnicity.[36]

Expect to Own in the Next Five to 10 Years

Results from the bivariate probit model governing whether current renters expect to own in the next five to ten years differ substantially from the current owner-occupancy model. Most striking, the race-related coefficients are all insignificant, indicating that we cannot reject the hypothesis that minority and white renters have similar expectations of future home purchase after controlling for the influence of credit barriers, demographic, and financial factors. Current marital status also has little influence on renter expectations, but older renters are significantly less likely to anticipate future homeownership, opposite from the impact of age on preferred current housing tenure. Presumably, this latter result reflects selection effects: older families with the strongest tendency to own would already have transited out of renting. Spouse working full time and expectations of future increase in income both have positive and significant impacts on renter expectations, whereas these variables have insignificant and negative effects in the preferred current housing tenure model. This likely indicates that the ability to accumulate future wealth positively affects renter expectations of future homeownership. Receiving an inheritance also has a sharply positive impact on renter expectations of owning, but central city/suburban status has little effect. The former result is similar to the preferred current tenure model, but the latter differs in that individuals living in less densely developed areas are more likely to prefer to own their current home. This difference could again reflect selection

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[35]Whether the head or spouse has been previously married was also included in the model because the dissolution of past marriages could reduce an individual's current wealth. Although these variables have some impact in the NotCC equation, they have no impact on housing tenure preferences.

[36]The "Other" race category includes people of Asian descent, Native Americans, and other less populous groups. Because it is not possible to separate these groups, results for the "Other" category are not discussed.

effects in that families with strong tastes for owning their current home may have already gravitated to the suburbs.

VI. The Impact of Borrowing Constraints on Owner-Occupancy Rates

In Figure 4, the first two columns report the sample share for which NotCC equals 0, labeled *Possibly Credit Constrained* in the figure, and the sample share for which NotCC equals 0 *and* families currently rent, labeled *Possibly Credit Constrained AND Renting*. For the full sample of owners plus renters, 14.2 percent of families currently rent and are possibly credit constrained. That estimate provides an upper bound on the impact of borrowing constraints on owner-occupancy rates.

The next three columns compare actual with predicted owner-occupancy rates that would prevail in the absence of borrowing constraints. Observe that eliminating borrowing constraints would have raised the 1998 owner-occupancy rate by 4.03 percentage points. Not surprisingly, borrowing constraints have the greatest impact on the homeownership rates of lower-income families: roughly 11 percentage points among families with total household income in the first decile, roughly 6.75 percentage points among families with income between the 10th and 50th percentiles, and little effect thereafter. Effects are also more pronounced among young (under age 35) and middle-aged families (between 35 and 55 years), raising homeownership rates for both groups by roughly 6.5 percentage points, with no effect on older families. Looking across racial and ethnic lines, homeownership rates rise most for Hispanics (6.7 percent) and least for African-American families (1.3 percentage points).

An additional perspective is obtained by comparing the predicted share of current owners across race. Observe that removing the influence of borrowing constraints does little to reduce the difference in owner-occupancy rates between white and nonwhite families. In contrast, in Figure 2, the difference in the frequency of married families between white and African-American households is 26.8 percentage points. Multiplying that value by the marital status partial in Figure 3 for the currently own regression yields 0.0555. Thus, if one were to raise the African-American marital status rate to that of the white population, *ceteris paribus*, the

African-American owner-occupancy rate would increase by 5.55 percentage points. This suggests that policymakers and industry officials seeking to narrow the homeownership gap between African-American and white households may need to look further than just the elimination of borrowing constraints.

Do credit barriers depress owner-occupancy rates by delaying access to owner-occupied housing or by permanently excluding some families from owning? The last three columns of figure 4 shed light on this question by examining renter expectations of future homeownership. Results indicate that the percentage of renters who expect to own in the next decade rises by 7.56 percentage points with the elimination of borrowing constraints. Assuming such expectations are realized, multiplying that figure by the current frequency of renters yields a 2.47 percentage point increase in the *eventual* attainment of owner-occupancy for the present total population of owners and renters (.0247 = .0756(1-.6738), where .6738 is the frequency of owners and 1-.6738 is the frequency of renters). In contrast, as noted above, elimination of borrowing constraints would increase owner-occupancy rates by 4 percentage points. Thus, the impact of borrowing constraints on renter expectations of future homeownership is small relative to the effect on the overall rate of homeownership. This suggests eliminating borrowing constraints accelerates the realization of renter expectations of owning a home. Equivalently, borrowing constraints depress current owner-occupancy rates in part by delaying access to homeownership as opposed to permanently excluding families from owning a home, consistent with results from Goodman and Nichols (1997).

Breaking out these last results by race yields a sharper picture. For whites, removing credit barriers increases renter expectations of future homeownership by 7.83 percentage points. This translates into an increase in the eventual attainment of homeownership among white families of 2.12 percentage points. For African American and Hispanic households, comparable calculations yield estimates of an increase in eventual attainment of homeownership of 3.28 and 5.18 percentage points, respectively. In contrast, the current housing tenure model predicts that with the removal of credit barriers, homeownership rates would increase 4.11, 1.31, and 6.72 percentage points for whites, African Americans, and Hispanics. Thus, although borrowing constraints appear to depress white homeownership rates at least in part by delaying renter access to homeownership (2.12 is well below 4.11), results for African Americans and Hispanics are more consistent with the view that credit barriers depress homeownership rates by permanently discouraging some renters from owning a home.

VII. Conclusions

Recent dramatic innovations in affordable mortgage lending along with the economic expansion of the 1990s have helped to raise U.S. homeownership rates to historic levels, from 64 percent in 1989 to 67 percent by 2000. Against that backdrop, this paper has sought to evaluate the extent to which the elimination of borrowing constraints would further expand opportunities for homeownership. This question is evaluated by applying discrete choice sample selection methods to unusually rich data from the 1998 Survey of Consumer Finances. A key feature of the SCF, upon which the estimation strategy is based, is that survey questions identify a group of households for whom borrowing constraints are not binding, *a priori*.

Is there opportunity for industry and government to expand homeownership through further relaxation of borrowing constraints? Results from the analysis suggest a qualified yes. If all borrowing constraints were suddenly removed, *ceteris paribus,* and all households could instantly change their housing tenure if they so chose, the owner-occupancy rate among nonfarm families in the United States would increase by just over 4 percentage points. Not surprisingly, these effects are distributed unequally across different subgroups of the population. Borrowing constraints have far more impact on homeownership rates among low-income families than any other group. Sizable opportunities to expand homeownership also exist among young and middle-aged families.

Further analysis suggests that for white households, borrowing constraints depress current owner-occupancy rates in part by delaying rather than by permanently excluding access to homeownership. That result is consistent with evidence from Goodman and Nichols (1997) and suggests that government and industry officials seeking to expand homeownership opportunities may want to consider mortgage products that alleviate borrowing constraints primarily in the early years of a mortgage. On the other hand, analogous estimates for African American and Hispanic families are suggestive that credit barriers depress owner-occupancy rates for these families by permanently discouraging some renters from homeownership. Because the approach used to identify these effects is indirect, these findings should be viewed with some caution. Nevertheless, the results here suggest that the degree to which borrowing constraints delay access to homeownership versus permanently exclude access to owning warrants further research.

Finally, model estimates also confirm that even in an environment free of borrowing constraints, household family and financial stability are very important determinants of whether families prefer to own: instability favors renting while stability favors homeownership. In addition, families in better health, and therefore more able to care for their properties, are more likely to prefer to own, as are higher wealth families who are more capable of absorbing financial risk. As an example of the potential impact of such effects on owner-occupancy rates, African-American marriage rates are roughly 25 percentage points below those of the white population. That difference alone is predicted to reduce the owner-occupancy rate of African-American families by roughly 5½ percentage points relative to the white population. Estimates such as these suggest that government and industry efforts to expand access to homeownership will be more fruitful if they include broad-based initiatives designed to enhance the social and financial stability of families, in addition to ensuring that affordable mortgage products are available.

APPENDIX

Figure 2 and 3 report the variable descriptions in each row to facilitate review. While most descriptions are complete, some require additional clarification. Under *Household head's race*, "Other" includes people of Asian decent, Native Americans, and other less populous groups. Because it is not possible to separate these groups into their own categories, results for the Other category are largely not discussed in the paper. Under *Stable income and employment*, "Know next yr income" equals 1 if the family usually knows what its income will be one year ahead. In the same category, "H # FT jobs > 1 yr" is the number of full time jobs lasting at least one year previously held by the head of household. Under *Inheritances, gifts, etc.*, "Received inherit/gift" equals 1 if the head of household or spouse has ever received an inheritance, major gift, or settlement. "$ value of inher/gift" is the dollar value of all inheritances, etc., received from 1995 on. "$ value expect inher/gift" is the dollar value of all inheritances, etc., that the household expects to receive in the future. Under *Central city/suburb,* "Ngh Bldg < 21 ft" equals 1 if the neighborhood buildings are within 21 feet of each other. A similar interpretation applies to the remaining variables in this category except "Density not known" for which the local density of development was not reported. That variable was included in the probit regressions, but results are not reported to conserve space. Under *Financial problems*, "H or S ever bankrupt" equals 1 if the head of household or spouse has ever filed for bankruptcy, while "Loan pays 2 mnth late" equals 1 if in the last year the family made any loan payments more than two months late.

Figure 1a: Percentage Race in the CPS and SCF for Various Years

	Farm Plus Nonfarm Populations								Nonfarm Only
	1989		1992		1995		1998		1998
	CPS	SCF	CPS	SCF	CPS	SCF	CPS	SCF	SCF
White	0.785	0.746	0.776	0.752	0.774	0.777	0.750	0.773	0.777
African American	0.112	0.129	0.114	0.127	0.116	0.128	0.119	0.121	0.119
Hispanic	0.077	0.080	0.081	0.075	0.082	0.057	0.094	0.074	0.072
Other	0.026	0.046	0.029	0.046	0.028	0.039	0.036	0.032	0.031

Figure 1b: Homeownership Rates in the CPS and SCF for Various Years

	Farm Plus Nonfarm Populations								Nonfarm Only
	1989		1992		1995		1998		1998
	CPS	SCF	CPS	SCF	CPS	SCF	CPS	SCF	SCF
White	0.693	0.703	0.695	0.701	0.709	0.705	0.723	0.717	0.73
African Am.	0.419	0.420	0.425	0.431	0.422	0.426	0.466	0.459	0.473
Hispanic	0.456	0.449	0.447	0.449	0.433	0.436	0.472	0.458	0.455
Other	0.506	0.536	0.520	0.542	0.505	0.517	0.535	0.540	0.558
Total	0.639	0.639	0.639	0.640	0.647	0.647	0.661	0.661	0.674

Figure 2: Sample Means of Probit Model Variables for the Full Sample and for Selected Subgroups

(All values are calculated with sampling weights to be representative of the United States)

	Full Sample	*White*	*African American*	*Hispanic*	*Other*	*Inc 0-10 Prcntl*	*Inc 10-25 Prcntl*	*Inc 25-50 Prcntl*	*Age < 35*	*Bld < 21 ft Apart*
Dependent Variables										
Own home	0.674	0.730	0.473	0.455	0.558	0.344	0.494	0.602	0.407	0.544
Rent but expect to own	0.050	0.045	0.158	0.074	0.093	0.092	0.061	0.069	0.082	0.071
"Not" credit constrained	0.717	0.750	0.551	0.630	0.715	0.668	0.720	0.662	0.511	0.663
Independent Variables										
Head's gender										
Male	0.722	0.744	0.517	0.800	0.773	0.416	0.467	0.675	0.744	0.678
Head's race										
White	0.777	1.000	-	-	-	0.582	0.708	0.749	0.691	0.678
African Amer.	0.119	-	1.000	-	-	0.297	0.167	0.121	0.136	0.166
Hispanic	0.072	-	-	1.000	-	0.081	0.101	0.108	0.127	0.116
Other	0.031	-	-	-	1.000	0.040	0.025	0.022	0.045	0.040
Head's marital status										
Married	0.525	0.551	0.283	0.588	0.653	0.172	0.266	0.420	0.449	0.435
Divorced	0.127	0.132	0.140	0.082	0.070	0.160	0.166	0.168	0.065	0.148
Head's age										
Age	48.841	50.199	46.240	40.861	43.369	49.278	53.074	49.190	27.964	46.720
Under 35 years	0.229	0.203	0.261	0.404	0.334	0.333	0.264	0.250	1.000	0.283
35 to 55 years	0.426	0.422	0.440	0.432	0.477	0.253	0.248	0.371	-	0.423
Over 55 years	0.345	0.375	0.300	0.164	0.189	0.413	0.488	0379	-	0.294
Size of household										
# in household	2.589	2.475	2.639	3.475	3.183	2.100	2.243	2.389	2.822	2.526
Current employment status										
Head works full time	0.637	0.634	0.565	0.733	0.759	0.216	0.368	0.623	0.804	0.645
Spouse works full time	0.284	0.289	0.219	0.276	0.416	0.070	0.084	0.191	0.322	0.251
Spouse works part time	0.076	0.081	0.032	0.085	0.083	0.038	0.037	0.060	0.082	0.064
Current, past, exp. inc										
Total family inc in 97	5.29E+04	5.85E+04	2.87E+04	3.17E+04	5.49E+04	4.28E+03	1.24E+04	2.48E+04	3.64E+04	4.31E+04
Real inc rose last 5 yr	0.203	0.208	0.184	0.189	0.182	0.095	0.063	0.140	0.293	0.207
Real inc exp to rise	0.234	0.213	0.295	0.333	0.298	0.256	0.202	0.207	0.412	0.264

Figure 2 (con't.): Sample Means of Probit Model Variables for the Full Sample and for Selected Subgroups

(All values are calculated with sampling weights to be representative of the United States)

	Full Sample	*White*	*African American*	*Hispanic*	*Other*	*Inc 0-10 Prcntl*	*Inc 10-25 Prcntl*	*Inc 25-50 Prcntl*	*Age < 35*	*Bld < 21 ft apart*
Stable inc and emp										
Know next yr income	0.722	0.757	0.586	0.570	0.729	0.536	0.568	0.709	0.614	0.685
H # of FT jobs > 1 yr	2.324	2.371	1.949	2.498	2.174	0.590	1.460	2.308	2.170	2.392
Health status										
Head bad health	0.051	0.046	0.086	0.052	0.057	0.138	0.109	0.051	0.009	0.060
Spouse bad health	0.021	0.025	0.008	0.018	0.000	0.020	0.032	0.027	0.010	0.025
Head's education										
Less than high school	0.187	0.149	0.270	0.474	0.161	0.445	0.373	0.206	0.162	0.187
High school	0.291	0.294	0.341	0.241	0.157	0.275	0.334	0.348	0.285	0.272
Some college	0.238	0.240	0.238	0.164	0.351	0.179	0.182	0.262	0.275	0.255
College degree	0.165	0.181	0.115	0.074	0.165	0.075	0.082	0.136	0.197	0.179
Graduate degree	0.119	0.136	0.036	0.047	0.166	0.025	0.030	0.048	0.081	0.108
Inheritances, gifts, etc.										
Received inherit/gift	0.203	0.237	0.109	0.044	0.077	0.119	0.172	0.208	0.121	0.165
$ value of inher/gift	8.06E+03	9.15E+03	5.34E+03	1.87E+03	5.32E+03	2.57E+03	3.64E+03	4.79E+03	4.87E+03	5.20E+03
$ value expect inher/gift	2.86E+04	3.29E+04	2.03E+03	8.46E+03	7.03E+04	4.94E+03	2.27E+04	1.35E+04	4.43E+04	2.46E+04
Central city/suburb										
Ngh bldg < 21 ft	0.464	0.405	0.647	0.746	0.592	0.534	0.506	0.523	0.574	1.000
Ngh bldg 21 to 100 ft	0.380	0.411	0.298	0.205	0.321	0.322	0.362	0.340	0.332	-
Ngh bldg > 100 ft	0.127	0.153	0.041	0.030	0.054	0.111	0.112	0.115	0.062	-
Density not known	0.029	0.032	0.014	0.019	0.033	0.033	0.020	0.023	0.032	-
Previous marriages										
Head prev married	0.175	0.191	0.136	0.106	0.080	0.147	0.125	0.172	0.043	0.159
Spouse prev married	0.120	0.133	0.069	0.067	0.109	0.030	0.052	0.102	0.049	0.091
Financial problems										
H or S ever bankrupt	0.084	0.086	0.102	0.046	0.069	0.065	0.078	0.091	0.065	0.091
Loan pays 2 mnth late	0.060	0.051	0.121	0.049	0.071	0.059	0.080	0.088	0.091	0.070

Figure 3: Three-Celled Bivariate Probit Model Estimates of Who Prefers to Own in the Absence of Borrowing Constraints[a]

	Preferred Current Housing Tenure *Sample Includes Owners and Renters*				**Expect to Own in the Next 5 to 10 Yrs** *Sample Includes Renters Only*			
	Not Credit Constrained		**Prefer to Own If Not Credit Constrained**		**Not Credit Constrained**		**Expect to Own If Not Credit Constrained**	
	Partial[b]	t-ratio	Partial[b]	t-ratio	Partial[b]	t-ratio	Partial[b]	t-ratio
Household head's gender								
Male	0.00619	0.346	0.03501	1.704	0.00472	0.163	0.05042	1.116
Household head's race								
African American	-0.09242	-4.899	-0.08034	-2.883	-0.06556	-2.14	0.05478	1.069
Hispanic	-0.00859	-0.362	-0.08830	-2.841	0.01504	0.397	-0.01226	-0.212
Other	-0.01454	-0.471	-0.11590	-3.108	0.02314	0.412	-0.02290	-0.294
Current marital status								
Married	0.06254	3.414	0.20706	8.336	0.01514	0.453	0.02782	0.553
Divorced	-0.03166	-1.555	0.06118	2.669	-0.02737	-0.761	0.07032	1.293
Household head's age								
Age if under 35 years	0.00766	4.698	0.00132	0.667	0.00567	1.886	-0.00915	-1.947
Age if 35 to 55 years	0.00781	7.475	0.00521	3.804	0.00559	2.842	-0.00726	-2.376
Age if over 55 years	0.00731	9.680	0.00468	4.676	0.00748	5.478	-0.00909	-3.868
Size of household								
# people in household	-0.01238	-2.589	0.02213	3.213	-0.01764	-2.001	0.00247	0.156
Current employment status								
Head works full time	-0.01839	-1.044	0.09961	4.347	-0.02714	-0.868	0.12318	2.624
Spouse works full time	-0.02766	-1.816	0.02438	1.165	-0.02736	-0.834	0.15079	3.189
Spouse works part time	-0.00189	-0.083	0.01020	0.320	-0.06885	-1.190	0.13940	1.633
Current, past, and Exp income								
Total household inc in 1997	2.12E-08	2.128	2.78E-09	0.485	0.00000	1.771	0.00000	-0.773
Real inc rose last 5 yrs	0.02190	1.476	0.03383	1.696	0.01213	0.383	-0.00114	-0.025
Real inc exp to rise next 5 yrs	-0.01174	-0.849	-0.03966	-2.059	-0.01379	-0.535	0.07507	1.948
Stability of income and emp.								
Usually know next year's inc	0.02628	2.069	0.05811	3.539	0.00826	0.349	0.06148	1.649
Head's # of FT jobs > 1 yr	-0.00802	-4.032	-0.00839	-2.728	-0.00431	-1.169	0.00005	0.010
Health status								
Head in bad health	-0.01792	-0.543	-0.04065	-1.247	-0.06602	-1.275	-0.05565	-0.504
Spouse in bad health	-0.14029	-3.234	-0.15615	-3.029	-0.09405	-1.114	0.10787	0.832
Household head's education								
Less than high school	-0.04920	-2.501	-0.08793	-3.755	-	-	-	-
Some college	-0.03143	-1.889	0.01313	0.611	-0.00200	-0.069	0.03656	0.817
College degree	0.00822	0.450	0.04483	1.979	-0.01398	-0.393	0.13902	2.533
Graduate degree	0.03248	1.609	0.04603	1.862	0.03270	0.389	0.12635	1.834

Figure 3 (con't).: Three-Celled Bivariate Probit Model Estimates of Who Prefers to Own in the Absence of Borrowing Constraints[a]

	Preferred Current Housing Tenure Sample Includes Owners and Renters				**Expect to Own in the Next 5 to 10 Yrs** *Sample Includes Renters Only*			
	Not Credit Constrained		***Prefer to Own If Not Credit Constrained***		***Not Credit Constrained***		***Expect to Own If Not Credit Constrained***	
	Partial[b]	**t-ratio**	**Partial[b]**	**t-ratio**	**Partial[b]**	**t-ratio**	**Partial[b]**	**t-ratio**
Recent & expected gifts, inheritances, settlements								
Rcd at least 1 since 80	-0.01092	-0.713	0.07099	3.584	-0.04535	-1.174	0.15652	2.749
$value of inherit, gift since 95	9.55E-09	0.448	9.48E-08	0.709	0.00000	-0.012	0.00000	0.467
$value of expect inheritance	3.52E-09	0.435	5.16E-08	1.330	0.00000	-0.392	0.00000	1.329
Central city/suburb status								
Ngh bldg 21 to 100 ft apart	0.03151	2.364	0.11531	6.649	-0.00028	-0.011	-0.0300	-0.727
Ngh bldg > 100 ft apart	0.06530	3.208	0.12327	5.023	0.11614	1.983	0.03488	0.429
Previous marriages								
Head previously married	-0.03289	-1.910	-0.00480	-0.222	-	-	-	-
Spouse previously married	-0.02333	-1.175	-0.00959	-0.348	-	-	-	-
Financial problems								
H/S ever file for bankruptcy	-0.20573	-10.013	0.00601	0.149	-0.22033	-4.888	-	-
Loan payments 2 mos. late	-0.21752	-8.548	-0.06431	-1.148	-0.18617	-4.548	-	-
Constant	-0.12303	-2.395	-0.36955	-4.558	-0.08667	-0.981	-0.09983	-0.512
$\sigma_{NotCC,Own}$			-0.3484	-1.807			-0.4521	-2.059
Total observations		4,142		4,142		1,189		1,189
Censored observations		0		984		0		505
Uncensored observations		4,142		3,158		1,189		684

[a]Controls for sample selection are based on who is not versus who may be credit constrained. Partial derivatives are presented to facilitate interpretation.

[b]Partial derivatives were calculated by forming $t_{partial} = t \cdot [\sum w_i \cdot f(x_i t)] / \sum w_i$,where t is the probit model coefficient for the tenure equation, $f(x_i t)$ is the unit normal density function, w_i is the sampling weight for observation i, and $\sum w_i$ is the sample size (appropriately weighted). See the text for additional details.

Figure 4: 1998 Actual and Predicted Owner-Occupancy Rates (%) [a,b]

	Upper Bound Impact of Credit Constraints Sample Includes Owners and Renters		*Preferred Current Housing Tenure Sample Includes Owners and Renters*			*Expect to Own in the Next 5 to 10 Yrs Sample Includes Renters Only*		
	Possibly Credit Constrained	*Possibly Credit Constrained AND Renting*	*Actual Percent Own*	*Predicted[c] Percent Own*	*Actual—Predicted Percent Own[c]*	*Actual Expect to Own*	*Predicted[c] Expect to Own*	*Actual—Predicted Expect to Own[c]*
Total	28.33	14.20	67.38	71.42	4.03	26.72	34.28	7.56
By Race								
White	24.99	11.20	72.95	77.07	4.11	27.25	35.08	7.83
African Amer.	44.86	27.18	47.31	48.62	1.31	26.01	32.24	6.23
Hispanic	36.96	24.26	45.49	52.22	6.72	23.48	32.99	9.51
Other	28.52	15.97	55.79	62.14	6.35	31.09	34.92	3.83
By Income Percentile								
0 to 10th	33.17	26.93	34.41	45.51	11.10	12.01	22.20	10.19
10th to 25th	28.02	19.08	49.43	56.55	7.13	12.36	25.52	13.16
25th to 50th	33.81	18.52	60.20	66.85	6.65	28.18	36.22	8.04
50th to 75th	32.81	13.24	74.29	77.25	2.97	46.08	46.44	0.03
75th to 90th	19.03	4.43	88.63	86.89	-1.7%	52.12	51.49	-0.63
90th to 100th	14.07	1.51	93.64	91.36	-2.28	57.63	59.47	1.84
By Age								
< 35 years	48.85	32.23	40.72	47.07	6.35	37.28	45.93	8.65
35 to 55 yrs	32.62	13.72	71.47	78.00	6.53	27.74	37.45	9.71
> 55 years	9.44	2.84	80.00	79.41	-0.59	4.90	5.82	1.63
By Distance Between Building								
< 20 feet	33.68	20.19	54.42	59.28	4.86	27.62	35.75	8.13
21 to 100 ft	25.56	10.24	77.43	80.72	3.28	25.46	32.85	7.39
> 100 ft	15.85	3.30	85.68	87.15	1.47	29.00	29.76	0.76

[a] All figures were weighted using the modified 1998 SCF weight x42001 to ensure that the values are representative of the United States for the respective subsample (see Kennickell [1999] for details).
[b] Sample excludes farms. Mobile home occupants are counted as owners if they own either the land, the unit, or both.
[c] Coefficients from the owner-occupancy/expected owner equations in figure 3 were used to calculate the predicted values.

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