**Liquidity constraints and housing prices:**
**Theory and evidence from the VA Mortgage Program**

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**Abstract**

This paper employs a simple intertemporal model to show that presence of liquidity constraints can depress the price of a durable good below its net present rental value, regardless of the overall supply elasticity. The existence of price effects implies that the relaxation of liquidity constraints is not Pareto improving, and may in fact be regressive. Historical evidence, which exploits the fact that a clearly identifiable group, war veterans, enjoyed the most favored access to mortgage credit in the postwar era, supports the model. The results suggest that more recent mortgage market innovations have served primarily to increase prices rather than home ownership rates, and that such innovations have the potential to exacerbate socioeconomic disparities in ownership rates. (JEL D91; E21; G21; R21)

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## I. Introduction

In the United States, few policy goals inspire less political controversy than the promotion of home ownership. Both Democratic and Republican administrations have cited increased home ownership rates as an important priority, and have pointed towards recent increases in this rate as an accomplishment. Since the Great Depression, government efforts to ease the path to home ownership have focused on generating greater and more widespread access to credit markets. Such efforts, described in detail in Section II below, concord with a view that liquidity constraints prevent many households from engaging in socially preferred behavior.

In recent economic literature, the role of liquidity constraints in altering behavior is a matter of some debate. A number of articles, both examining housing tenure (Duca and Rosenthal 1994; Gyourko et al. 1999; Haurin et al. 1997; Linneman et al. 1997; Linneman and Wachter 1989; Listokin et al. 2001; Rosenthal, 2002; Zorn 1989) and more general consumption patterns (Hayashi, 1985; Chah, Ramey, and Starr, 1995; Stepens, 2002) support the hypothesis that consumer behavior changes in the presence of borrowing constraints. Borrowing constraints have frequently been implicated as an explanation for empirical failures of the life-cycle and permanent income hypotheses (see Deaton, 1992, for a thorough review). On the other hand, several prominent papers have challenged the notion that liquidity constraints strongly influence behavior (Carneiro and Heckman 2002; Cameron and Taber 2004; Hurst and Lusardi 2004). Households facing liquidity constraints differ in many respects from unconstrained households, and it is these other differences that may truly explain differences in observed behavior. If this is the case, then relaxing liquidity constraints would have little or no impact on equilibrium consumption or investment patterns.

This paper focuses on a second reason why the relaxation of liquidity constraints may have

limited impact on behavior: because greater access to credit has the potential to raise the equilibrium price of durable goods. Price effects may explain, for example, why numerous mortgage market innovations over the past twenty years have had only a negligible impact on home ownership rates. The possibility that liquidity constraints depress prices has been raised in previous theoretical literature (Ranney 1981; Stein 1995; Ortalo-Magne and Rady 2002). Empirical studies, however, often ignore the potential for price effects or assume that any such effects are negligible.

Does the relaxation of borrowing constraints lead to higher prices for assets and durable goods? Is this effect large enough to be empirically relevant? Existing literature is largely mute on these questions; this paper provides answers to each of them.

This study extends the existing theoretical literature on liquidity constraints and prices by debunking the primary argument of those who claim that price effects are negligible: that highly elastic long-run supply curves limit the potential for demand increases to translate into price increases. A very simple intertemporal model, inspired by existing models of tenure choice in the housing market (Artle and Varaiya, 1978; Brueckner, 1986; Henderson and Ioannides, 1983; Stein 1995), shows that borrowing constraint relaxation can make asset values rise relative to the present discounted value of rents even when supply is perfectly elastic.

This study also presents an empirical analysis of liquidity constraints and prices that examines both time-series evidence on the relation of housing values to rents and cross-sectional variation generated by a large government program that selectively relaxed borrowing constraints for a clearly identifiable subset of the population: the VA mortgage program. The evidence suggests that mortgage market innovations contributed to a remarkable rise in housing values relative to rents between 1940 and the present. In particular, the ratio of housing values to rents displays sensitivity

to the share of veterans in the market in the time period when the VA mortgage program had the greatest impact on home ownership rates. Section II below provides basic information on the evolution of housing prices in this era; following the theoretical discussion in section III, section IV presents time-series evidence and section V analyzes the implications of the VA mortgage program.

The existence of price effects implies that the relaxation of borrowing constraints, though welfare improving by Hicks-Kaldor criteria, is not Pareto improving. If easier credit creates winners and losers, how might we most accurately characterize consumers who gain and those who suffer? While section III discusses this question in theory, section VI presents empirical evidence to suggest that households of low socioeconomic status benefit least from innovations in the credit market. Given the billions of dollars spent on improving access to mortgage markets and otherwise subsidizing home ownership in the United States, the distributional implications of promoting easy access to credit should be widely noted.

## II. Motivating evidence: housing prices, the mortgage market, and home ownership

The mid-twentieth century witnessed a profound change in both home ownership patterns and the relationship between asset prices and rental rates in the housing market.[1] As Table 1 indicates, the home ownership rate in the U.S. increased by nearly 50% between 1940 and 1960. After 1960, ownership rates continued to increase, but at a markedly slower rate. Table 1 also shows

[1] Economists may be skeptical of the importance of home ownership as an indicator of well-being, relative to measures such as household income or consumption. Interest in home ownership can be justified on the grounds that consumers are myopic and need to employ commitment devices in order to accumulate assets (Laibson 1997). There is also evidence that housing tenure decisions have nonmarket implications for child and family development (Boehm and Schlottmann, 1997; Green and White 1997) and citizenship (DiPasquale and Glaeser 1997; see Rohe et al. 2002 for a literature review). Finally, public economists take interest in home ownership because of the income tax treatment of mortgage payments and other policy initiatives (Engelhardt, 1997; Follain et al. 1987; Poterba, 1992; Poterba 1984; see Hendershott and White 2000 for a literature review). See Collins and Margo (2001) and Masnick (2001) for analysis of home ownership patterns in the U.S. across the twentieth century.

that this marked increase in home ownership was accompanied by a substantial increase in the price of the typical owner occupied housing unit, relative to the cost of the typical rental unit. In 1940, the median owner-occupied housing unit was worth roughly 140 times median monthly rent. By 1960, this ratio had increased to at least 180.[2] The ratio has remained at or above this general level since, with some evidence of further increases in recent decades.[3]

This escalation, while impressive, may reflect changes in the physical characteristics of owner- and renter-occupied units over time. A measure of constant-quality owner-occupied housing prices can be obtained from the National Income and Product Accounts beginning in 1929. Figure 1 plots this price index, converted to a logarithmic scale, along with the corresponding price index for all personal consumption expenditures. Both indices have been normalized to a value of zero in 1947. This graph shows that over the postwar period, constant-quality housing values have tracked overall price levels fairly closely, with two noteworthy exceptions. Housing prices accelerated above inflation in the immediate postwar era and remained relatively higher than the prices of other consumer goods until the mid-1970s.[4] This pattern corroborates the acceleration of owner-occupied housing values displayed in Table 1. In the late 1970s, housing prices tracked general price levels quite closely through a period of high overall inflation. Beginning in 1982,

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[2] It is not possible to pinpoint the exact ratio after 1940 because housing value data are intervalled. Census data on housing prices and rents are not available prior to 1940.

[3] The sample of owner-occupied units used to construct these statistics consists of nonfarm, non-condominium occupied single family units with no commercial uses, on lots of less than ten acres. In 1940, it is not possible to distinguish multifamily units or units with commercial units, however respondents were instructed in that year to omit the value of units they did not occupy or commercial space when providing their estimate of market value.

[4] Figure 1 also shows that the price of owner-occupied housing fell relative to the price of other consumer goods during World War II. This may reflect the wartime rationing of goods other than owner-occupied housing.

housing prices began another period of acceleration.[5] This acceleration has persisted to the present day.

These two periods of housing price escalation coincide with major relaxations in borrowing constraints, as seen in the timeline of major mortgage market innovations presented in Table 2 (Jackson, 1985; Bruskin et al. 2000; Listokin et al. 2001; Martinez 2000). At the onset of the Great Depression, when the time series graph in Figure 1 begins, families interested in purchasing a home generally were expected to make a down payment of at least 30% of the home's value. The loan they received would not be fully amortized, requiring a balloon payment (or refinancing) after five to ten years. During the depression, a large number of households defaulted on their mortgage payments, leading to foreclosure in many cases. Over an eleven-year period beginning in 1933, government innovations, including the creation of the Federal Housing Administration (FHA) and Veterans' Administration (VA) mortgage insurance programs, and the establishment of a secondary market for mortgage debt, radically changed the nature of the mortgage market. By 1944, it was possible for many households to receive a fully amortized, twenty to thirty-year mortgage with down payments of 10% or less. The VA mortgage program was the most generous of those introduced during this time period, since it enabled eligible veterans to purchase a home with no down payment whatsoever and no mortgage insurance premium. As with the FHA program, the Federal government insured lenders against default risk.

The impact of these innovations on home ownership rates and housing prices was somewhat delayed by depression and war. Ownership rates reached a local minimum in the 1940 Census, and

[5] Interestingly, these time series indicate that the period of rapid home price acceleration in the 1970s, analyzed by Mankiw and Weil (1989) and many subsequent researchers, can be viewed as a simple manifestation of overall price level increases or increases in housing unit quality over the same time period.

Figure 1 shows that housing prices continued to sink relative to other consumer goods until the end of World War II. The coincidence of rapidly increasing ownership rates and housing prices in the postwar era suggests an important role for relaxed liquidity constraints.

The second period of escalating relative house prices, beginning around 1982, coincides with further innovations in the mortgage market. Until the 1980s the secondary mortgage market, which encouraged lending by pooling the risk that individual financial institutions assumed when extending credit, was an avenue open only for loans that "conformed" to underwriting standards governing borrower creditworthiness and property characteristics. A secondary market for "non-conforming" loans, mortgages that entities such as the Federal National Mortgage Association (FNMA) refused to securitize, grew rapidly through the 1980s and into the 1990s. The definition of "conforming" loans was itself relaxed in response to Federal legislation in 1992. By 1994, lenders had initialized programs that allowed qualified households to borrow more than the value of a home, effectively creating a negative down payment that could be applied towards closing costs or received in cash. These innovations enabled some previously ineligible households to purchase a home, and provided many others with increased buying power given current wealth levels (Bruskin et al., 2000). As Table 1 indicates, these new mortgage market innovations are not associated with any appreciable change in home ownership rates. They do, however, coincide with the current period of relatively high prices for owner-occupied housing, as shown in Figure 1.

Could factors other than credit market innovations explain these broad trends? Fluctuations in real interest rates undoubtedly influence the ratio of housing prices to rents; however the trends identified in this section are of too broad a period to be explained by cyclical variation. Changes in the value of the tax subsidy for mortgage interest, which has been in place since 1913, might also

be expected to influence housing prices (Poterba, 1984). Note, however, that housing prices failed to accelerate ahead of inflation during the "bracket creep" era of the late 1970s and 1980s. Prices did accelerate, however, during the latter part of the 1980s, a time of broad declines in marginal tax rates.

## III. Tenure choice, liquidity constraints, and prices: a simple model

There have been several noteworthy theoretical studies of housing prices and more generally of housing demand (Ranney, 1981; Poterba, 1984; Schwab 1982; Stein, 1995; Ortalo-Magne and Rady 2002; see Smith et al. 1988 for a review of housing market models). Stein (1995), in particular, uses a static model to show that borrowing constraints reduce housing prices. Stein's result can be attributed to the assumption of a fixed housing stock, and the fact that households must pay for housing prior to consuming and earning income. Under different assumptions, particularly if housing supply is close to perfectly elastic, this type of model produces radically different results.

The model presented here is an advancement over Stein's model in several respects. It is a two-period intertemporal model. There is no restriction on the price elasticity of housing supply. Finally, there is a fully integrated rental market for housing as well as a market for owner-occupied housing.[6] The model shows that liquidity constraints can depress the price of owner-occupied housing below the present value of rents, regardless of the local price elasticity of supply.

Agents $i$ derive utility from two consumption goods, housing services $H_i$ and a numeraire commodity $X_i$. They are endowed with varying amounts of a single physical asset, $A_i$, which can be

[6] In an appendix, Stein (1995) incorporates a rental market, but requires that renters pay an arbirtrary premium relative to owner-occupied housing prices. This model can also generate such a premium as a result of liquidity constraints, rather than by assumption.

interpreted as land. The overall amount of land in the economy is fixed. They are also endowed with varying amounts of effective labor, $L_i$, which they supply inelastically to firms. Competitive firms purchase labor and rent land in competitive factor markets and transform them into the numeraire commodity. Any quantities of land not used in the production process in a given period can be costlessly transformed into housing services. In the first period, agents may buy, sell or rent land. Only consumers may purchase land, but either consumers or firms may rent it. Agents' net housing consumption in each period is $A_i$, less the net amount sold $A_i^S$ and the net amount rented to other agents $A_i^R$.

Using $X_{ti}$ to denote numeraire consumption in period $t$, assuming that utility is time-separable, and that agents making decisions in the first period discount second-period utility by a factor $\delta$, the consumer's maximization problem can be stated as follows:

(1) $\underset{X_{1i}, X_{2i}, A_i^S, A_i^R}{Max} \; U(H_i, X_{1i}) + \delta U(H_i, X_{2i})$,

subject to the accounting identity

(2) $H_i = A_i - A_i^S - A_i^R$.

Assuming that agents can store the numeraire commodity without risk or return between periods, the maximization in equation (1) is undertaken subject to the lifetime budget constraint:

(3) $VA_i^S + 2rA_i^R + 2L_iw - X_{1i} - X_{2i} \geq 0$,

where $w$ is the market wage, equal to the marginal product of effective labor, $r$ equals the marginal product of land in the production of the numeraire, and $V$ is the endogenously determined price of a unit of land.

Agents with asset holdings that are sufficiently small relative to their labor income will find it desirable to borrow against their second period income. Such borrowing is governed by the

relation:

(4) $VA_i^S + rA_i^R + L_i w - X_{1i} \geq -\mu_i$.

In words, excess of first-period consumption over total first-period income may not exceed $\mu_i$.[7] In scenarios where borrowing is prohibited, $\mu_i$=0. Less stringent regulations on borrowing involve some greater value of $\mu_i$.

With this setup, it is quite simple to show that per-unit housing values $V$ can fall below the present value of rents, $2r$, in the presence of liquidity constraints. Denoting the Lagrange multiplier from the lifetime constraint (3) as $\lambda_1$ and the multiplier from the intertemporal constraint (4) as $\lambda_2$, the first order conditions of the consumer's problem yield the following expression of a household's willingness to pay to own rather than rent a marginal unit of land, $V^*$, conditional on equilibrium rent levels $r$:

(5) $V^* = \frac{2\lambda_1 + \lambda_2}{\lambda_1 + \lambda_2} r$.

When the intertemporal constraint is nonbinding, this expression for $V^*$ reduces to $2r$, the present value of rents.[8] As liquidity constraints become more binding, increasing the shadow value of first-period assets, this expression falls further below the present value of rents.[9] Liquidity constrained consumers thus "borrow" against their asset holdings to increase their first period consumption.

To the extent that a sufficient number of unconstrained consumers exist in the economy, exchange can take place at a market price of $V$=$2r$. Constrained consumers thus gain some surplus from selling their asset holdings at a price higher than their willingness to accept. If the marginal

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[7] Stiglitz and Weiss (1981) present a model justifying the existence of credit rationing in markets with imperfect information.

[8] It can be shown that at least one consumer will always be unconstrained in this framework.

[9]

net purchaser of land, however, faces liquidity constraints, the equilibrium price must fall below the net present value of rents.[10]

The distributional implications of relaxing liquidity constraints are easily derived. Consumers can be divided into two groups on the basis of their value of $\lambda_2$. Initially, less-constrained households are net purchasers of housing, and realize the benefits of depressed prices either by increasing their housing consumption or renting additional units of land to finance numeraire consumption. More-constrained households adopt the opposite strategy, selling to finance additional numeraire consumption in the first period. Any general relaxation of liquidity constraints naturally yields positive benefits to net sellers of housing and negative benefits to net purchasers.

In this model, less-constrained households will tend to be those with high asset holdings relative to income, or equivalently low income in relation to asset holdings. It is thus difficult to determine a priori whether the relaxation of liquidity constraints is progressive or regressive. If households with low income also tend to have a low ratio of income-to-assets, then easier access to credit can actually be a regressive policy.[11]

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[10] Note that this model can thus endogenously generate premiums for rental housing, observationally equivalent to discounts for owner-occupied housing, similar to those derived by Henderson and Ioannides (1983) and imposed by assumption in Stein (1995). In practice, equilibrium in markets where the marginal net purchaser is liquidity constrained could be achieved either by a decrease in $V$ or an increase in $r$. Under the assumption of diminishing marginal product, higher $r$ implies greater labor intensity and lower equilibrium wages. With fixed asset holdings, a household with lower labor income is more likely to be credit constrained.

In theory, the guaranteed existence of at least one unconstrained consumer in the economy implies that the equality $V=2r$ could be maintained by simply having that consumer purchase any marginal units of land. Adding realistic complications to the model, such as monitoring costs associated with purchasing land in a market where one does not live, or asymmetric information associated with shocks to land productivity, could eliminate this implication.

[11] There is some empirical evidence to support the notion that credit access programs do not benefit the long-term poor. Goodman and Nichols (1997) show evidence that the FHA mortgage program primarily benefits young households – likely those with low asset holdings in relation to present and future income – rather than poor households.

Another implication of this model, which will be of importance for the empirical tests below, is that exogenous increases in the rental rate $r$ should increase the discount applied to $V$. Intuitively, since the net rental holdings of households are negative, a higher $r$ increases per-period income relative to net assets. Any increase in per-period income relative to asset holdings will increase households' desire to borrow to finance first-period consumption.

It is worth reiterating that this model's result holds without any reference to the price elasticity of supply for housing. In this model, the elasticity of housing supply depends on the second derivative of firms' production functions with respect to the physical asset. If the marginal product of land is constant, then the elasticity of supply for housing is infinite: increases in demand for housing translate entirely into changes in the quantity consumed.[12] Even when the rental price is constant, however, liquidity constraints can influence the purchase price of housing. Some previous empirical studies of liquidity constraints and tenure choice dismiss the possibility of price effects with arguments related to the supply elasticity of housing (see, for example, Monroe, 2001).[13] The results presented in this section show that these arguments are insufficient to assuage concerns that easier access to credit creates inflationary pressure in the housing market. The following sections present empirical evidence documenting such effects.

## IV. Time Series Evidence

As shown in Table 1, the ratio of median value to median rent has increased considerably

[12] In this model, the housing supply elasticity can only be infinite up to a point – determined by the total endowments of land in the economy. Also note that increased housing demand may have a feedback effect, since reduced utilization of land may change the marginal product of labor in production, and hence wages.

[13] Several empirical studies have concluded that the price elasticity of housing supply is considerably less than infinite (Capozza, Green and Hendershott 1996; Topel and Rosen 1988; Malpezzi and Mayo 1997).

over the past sixty years. Table 3 presents the results of simple tests of the hypothesis that this phenomenon reflects a gradual elimination of discounts for owner-occupied housing in high-rent housing markets. As illustrated above, such a pattern would be expected if consumers in these markets experienced a reduction in liquidity constraints.

These regressions make use of Integrated Public Use Microdata Sample (IPUMS) data collected from a series of U.S. Census enumerations beginning in 1940. The unit of observation in each specification is an owner-occupied housing unit.[14] The dependent variable in each specification is the logarithm of the owner's self-reported estimate of the unit's market value.[15] The independent variable of primary interest is the logarithm of median rent in the housing market.[16] In this and all subsequent analyses, metropolitan statistical areas (MSAs) will serve as a proxy for a housing market area.[17]

The estimation is hampered to some extent by the absence of data on housing structural characteristics in the 1940 Census. For purposes of comparing results in 1940 and later years, Table 3 reports the results of 1970 specifications constrained to match the data limitations in 1940. Simple

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[14] Multifamily dwellings are excluded for all sample years except 1940. In 1940, owners of multifamily units were asked to estimate the value of their own unit only. It is not possible to determine which housing units are part of multifamily structures in these data. Farms are also excluded from the analysis.

[15] Previous research has shown that owners tend to overestimate their property's true market value, but estimation errors appear to be random (Goodman and Ittner, 1992). The 1940 value data are coded at the dollar level; the 1970 value data are intervalled.

[16] Since the independent variable of interest is a sample statistic, each observation in the regression is weighted by the square root of the sample size used in calculating the statistic, in order to avoid heteroskedasticity problems. Unweighted regressions yield similar results, though ln(rent) coefficients in both 1940 and 1970 are smaller in absolute value, as should be expected when noisier observations are given greater weight in the regression.

[17] Although the unit of observation in these regressions is an owner-occupied housing unit, the number of independent observations is effectively restricted to the number of MSAs in each year. For this reason, the Huber-White correction has been applied to the standard errors in these regressions.

univariate regressions of log housing values on log median rents, reported in the table's first and third columns, show a relationship that increases in magnitude between 1940 and 1970.[18] The point estimate in the first column suggests that metropolitan areas with higher rents in 1940 had less-than-proportionately higher values. The 1970 point estimate suggests the reverse, that ratios of value to rent were higher in high-rent markets.[19] While suggestive, it should be noted that neither of these coefficients can be statistically distinguished from one, and they are not significantly different from each other.

The ideal source of variation in median rents for this exercise would be exogenous differences across markets in terms of local amenities or opportunity costs of supplying housing units. The median rent data used here might confound this type of variation with that arising from characteristics of a metropolitan area's residents, most notably their income levels. To more effectively isolate location-based variation in median rents, the remaining regressions in Table 3 control for a basic set of housing market characteristics, including householders' median age, the fraction of householders working in manufacturing industries, the fraction of householders with a college education, and Census region effects. The 1970 regression also includes a measure of recent housing supply growth, the fraction of all housing units built after 1960, to account for the possibility that supply responses may have differentially closed gaps in rents and values across housing

[18] In an infinite-horizon model without liquidity constraints, $V = r/\delta$. Taking logs, we find $\ln(V) = \ln(r) - \ln(\delta)$. The intercept of this regression can thus be interpreted as the negative value of the logarithm of the rate of return used to compute housing values. The parameter estimates in Table 4 yield annual rates of return on the order of 6.8 to 7.7 percent. One alternative interpretation of the slope coefficient in the 1940 regression is that the relevant rate of return on investments in high rent markets is higher.

[19] In 1970, about 4% of owner-occupied housing units have topcoded value observations. This topcoding most likely biases the rent coefficient downward; omission of topcoded observations from the sample results in lower coefficient estimates.

markets.[20]

Controlling for housing market characteristics, the relationship between rents and values across housing markets in 1940 weakens considerably. For every 1 percent increase in median rents, predicted owner-occupied housing values increase by roughly 0.3 percent. A very different pattern emerges in 1970: the effect of a 1 percent increase in median rent is statistically indistiguishable from a 1 percent increase in values. The 1940 coefficient is significantly less than one and significantly different from the 1970 coefficient. This pattern corresponds exactly to the model's predictions: in a liquidity-constrained world, housing values in high-rent areas are discounted; when the constraints are relaxed these discounts disappear.[21]

The ideal source of variation in median rents would also account for differences in the quality of renter- and owner-occupied housing units across markets. Unfortunately, it is impossible to account for these factors using 1940 IPUMS data, since they contain no information on housing unit structural characteristics. In 1970, however, it is possible to perform such an exercise; the fifth regression in Table 3 presents the results. In this regression, the median rent variable for each metropolitan area has been replaced with each area's mean residual from a regression of ln(median

[20]The 1940 IPUMS data lack information necessary to construct this data; omission of this variable from the 1970 regressions does not alter the results.

[21] An alternative explanation for the patterns in Table 3 focuses on cyclical effects. Since housing values should reflect the present discounted value of rents, rents should display more cyclical volatility than values. In recessions, the ratio of values to rents should be higher; in expansionary periods the ratio should be lower. According to the National Bureau of Economic Research, the 1940 Census was taken nearly two years into a seven year-long expansion, while the 1970 Census was taken in the midst of a nearly year-long recession. Note, however, that a universal reduction in the ratio of values to rents should alter the intercept term of these regressions, not the coefficient on median rent. The median rent coefficient would change only if recessions differentially affect the ratio of values to rents in high-rent markets.

rent) on a basic set of structural characteristics.[22] This modified rent variable thus accounts for differences in observed rental unit characteristics across markets. The same set of structural characteristics also appear as regressors in the reported specification, though their coefficients are not reported here. Controlling for observable differences in housing unit quality has a negligible effect on the estimated relationship between median rents and housing values. Comparing the fourth and fifth regressions in Table 4, the coefficient falls from 1.17 to 1.10, a difference that is not statistically significant. On the basis of this evidence, the failure to consistently account for housing unit quality across markets over time does not appear to influence the basic message reported in Table 4: that the ratio of housing values to median rents leveled upwards between 1940 and 1970.

The remaining three regression specifications reported in Table 3 extend the analysis forward from 1970 to 2000. Each specification replicates the final 1970 model, which adjust median rents in each market and controls for structural characteristics, for IPUMS samples derived from 1980, 1990 and 2000 Census data. As is true in 1970 data, the estimated effect on housing values of a one percent increase in median rent in 1980 is statistically indistinguishable from one percent. In 1990 and 2000, estimated coefficients exceed one by a significant margin. From one perspective, these larger coefficient estimates coincide with the second, post-1980 period of mortgage market innovation that further relaxed borrowing constraints for a broad segment of the population. The results imply, however, that the ratio of value to rents is now higher in high-rent cities, a pattern that cannot be explained with the simple model outlined in the preceding section. There are several

[22] Added structural characteristics include categorical variables for number of rooms, year structure built, presence of basement, number of units in structure (rental units only), whether unit is detached (owner-occupied units only), number of units at address (rental units only), number of bathrooms, and presence of window or central air conditioning.

possible explanations for this finding. Higher relative values in high rent markets could reflect divergent expectations about future rent levels: consumers may now expect the fastest rent increases in the markets with highest current rent. Alternatively, borrowing constraints may now pose the greatest limitation for consumers in low rent markets. Finally, unobserved components of housing quality could be increasing most rapidly in the owner-occupied housing stock within high-rent MSAs.

Overall, the time series evidence can be regarded as suggestive but far from definitive. The escalation of owner-occupied housing values in high rent markets is consistent with the pattern expected with relaxed borrowing constraints, but it is clear that access to credit cannot by itself explain all of the evidence shown in Table 3.

## V. Repeated cross-sectional evidence: the VA mortgage program

### *A. Did the VA mortgage program influence housing tenure?*

The empirical tests of the relationship between liquidity constraints and prices presented below rely on the presumption that a readily identifiable group, war veterans, had significantly greater access to credit than otherwise identical householders in 1970. Given the plethora of mortgage market innovations that drove the postwar boom in home ownership, it is not inherently clear that the advantaged conferred by the VA mortgage program was significant. The probit specifications presented in Table 4 address this concern by looking for evidence that veterans' greater access to credit translated into higher probabilities of home ownership.

These ownership probits use data on householders derived from the same IPUMS samples used in the preceding section. Coefficients reported in the table indicate the marginal effect of a one-

unit change in the independent variable when all other variables are set equal to their respective means.[23]

Table 4 first addresses the concern that veteran status might be a poor proxy variable for access to credit. Such a concern would be valid if veterans had greater (or lesser) demand for owner-occupied housing for other unobserved reasons. To test the validity of these concerns, the first specification examines home ownership in 1940, prior to the implementation of the VA mortgage program. While ownership is significantly more common among households headed by older, married, male, non-black, and native-born individuals, and varies significantly by household size and region, there is no evidence to suggest that veterans were either more or less likely to own a home than otherwise identical householders.

A comparable probit specification for 1970, reported in the second column, distinguishes veterans by their projected eligibility for the VA mortgage program.[24] Ineligible veterans, similar to all veterans in 1940, have home ownership patterns statistically indistinguishable from otherwise identical householders. Eligible veterans, on the other hand, are significantly more likely to own their home. When all other variables are evaluated at their means, the magnitude of the veteran effect is 7.6 percentage points. Eligibility for the VA mortgage program thus confers an advantage comparable to a one-standard deviation increase in wage and salary income, or two additional years

[23] Summary statistics for all reported regression covariates appear in Appendix Table A1.

[24] Only veterans who served after 1941 are eligible for the VA mortgage program. The IPUMS data does not explicitly list dates of service, but rather identifies whether veterans served in a particular war (World War I, World War II, Korea or Vietnam) or in peacetime only. I count all veterans with service in World War II, Korea or Vietnam as eligible for the program. Moreover, I also include veterans serving in peacetime only who were under the age of 50 in 1970. The age distribution of peacetime-only veterans in the 1970 IPUMS sample is bimodal, with a local minimum occurring at age 50. In 1970, the VA mortgage benefit was also restricted to veterans who had served for at least 90 days. Unfortunately, the IPUMS data do not indicate a veteran's length of service. Thus the measure employed here likely misidentifies some ineligible veterans as eligible.

of age. Given that 43% of householders were eligible veterans in 1970, these results suggest that about 20% of the overall increase in home ownership rates between 1940 and 1970 can be attributed to the VA mortgage program.[25]

The final three specifications in Table 4 replicate the 1970 specification using later IPUMS samples. Eligibility for the VA mortgage program continues to exert a significant positive influence on the probability of home ownership in 1980, though the point estimate is only slightly more than half the size of the 1970 coefficient. In 1990 and 2000, eligibility for a VA mortgage ceases to be a significant predictor of home ownership. Given the degree of innovation in mortgage instruments beginning in the 1980s, this pattern is not surprising. This evidence suggests that veteran status can be used as a proxy for easier access to credit within a relatively narrow time window in the post-World War II era.

*B. Using veteran status as a proxy for access to credit*

Across housing markets, the fraction of households headed by an individual eligible for the VA mortgage program varied significantly, from approximately one-fourth to one-half in 1970. The regression results presented in Table 5 use the density of veterans in a metropolitan area as a proxy for the extent of borrowing constraint relaxation in that area. Theoretically, the escalation of value-to-rent ratios in high rent markets should be more acute among markets with a high density of veterans. The independent variable of interest in these regressions is therefore the interaction

[25] A pooled specification, which allows a difference-in-difference style estimation of the impact of VA mortgage program eligibility, shows a statistically significant 13.5 percentage point marginal effect. As noted in Collins and Margo (2001), caution should be used in interpreting this result since the estimated coefficients on many householder characteristics change considerably across the 1940 and 1970 samples.

between veteran density and the logarithm of median rent.[26] Each regression also includes the main effects of median rent and veteran density, as well as the set of housing market characteristic controls introduced in Table 3.[27] The central assumption underlying this strategy is that veteran status does not correlate with the unobserved component of demand for owner-occupied housing. The best available evidence to support this assumption appears in Table 4, which shows that veteran status is associated with higher home ownership rates only in the years when the VA mortgage program presented a distinct advantage over other readily available credit programs. Table 6 below will provide further evidence in support of this assumption.

The first regression reported in Table 5 predicts housing values as a function solely of housing market characteristics, including median rent and veteran density but excluding the interaction between these variables. The impact of a one percent increase in median rents on housing values is statistically indistinguishable from a one percent increase. Markets with a higher share of veterans feature higher housing values, other things equal. Among the eight other market characteristics controlled for in this specification, only householder education levels and location in the Southern Census region have a significant impact on housing values. Owner-occupied housing costs more in relatively educated markets, and less in the South.

This first regression supports the hypothesis that veterans, by virtue of their superior access to credit, have the potential to prop up housing values in markets where they form a sizable group.

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[26] As in Table 3, the unit of observation is the owner-occupied housing unit. Standard errors have been adjusted to reflect clustering of these observations by metropolitan area. Only veterans eligible for the VA mortgage program are used in calculating the veterans' share variable for 1970. Using the procedure for predicting eligibility described above, over 97% of veteran heads of household were eligible for the mortgage program in 1970.

[27] Exclusion of the housing market characteristic variables changes the magnitudes of some coefficients in Table 5, but does not influence their statistical significance.

However, this result also supports the simpler hypothesis that veterans simply consume more or better housing. The remaining regressions in Table 5 evaluate which explanation is most plausible.

The second specification introduces the interaction term between veteran density and the logarithm of median rent. The presence of householders with easier access to credit should matter less in low-rent markets, where few consumers were likely to be liquidity constrained in the first place. Consistent with this prediction, the interaction between veteran density and median rent is significant and positive, while the two main effects are significant and negative. The results imply that the relationship between median rents and housing values is weakest in markets with a relatively low veteran density; point estimates suggest that discounts for owner-occupied housing in high rent markets disappear once the veteran share exceeds roughly 41%. The point estimates also imply that veteran density predicts higher absolute housing values only in markets with median rents close to the observed maximum value in 1970. Introducing the interaction term also increases the magnitude and significance of several other covariates in the regression: the Western Census region now appears pricier than all others, areas with faster recent housing supply growth tend to have lower prices, and prices tend to be higher in manufacturing-oriented metropolitan areas.

The third and fourth columns in Table 5 present the results of specifications that incorporate controls for structural characteristics. As in Table 3, structural characteristics are both incorporated as explanatory variables and used to transform the median rent variable into a constant-quality measure of rental unit prices. These controls improve the fit of the model considerably, raising the $R^2$ measures from about 0.16 to 0.54.

With structural controls, and omitting the interaction between veteran density and median rent, there now appears to be a negative relationship between veteran share and housing values. The

positive coefficient observed in the table's first specification must, therefore, reflect the tendency for veterans to gravitate towards markets with larger or higher quality housing units. The significant negative coefficient assuages concerns that veterans are concentrated in markets with housing units of higher *unobserved* quality. Rather, it now appears that veterans for the most part congregate in markets where the price of owner-occupied housing is relatively low. This finding is consistent with the results in the preceding column.

The final regression in Table 5 re-introduces the interaction between veteran density and median rent. Once again, the interaction term is positive and significant while the two main effects are negative and significant. As in the table's second specification, the link between median rents and housing values is strongest in markets with a high veteran density. Unlike that prior specification, higher veteran density now predicts higher housing values in all markets, rather than just those with higher rents.

This reduced-form analysis cannot directly provide universal estimates of the elasticity of housing prices with respect to borrowing constraint regulations. The results can, however, gauge the impact of extending VA-style access to credit markets to an additional fraction of the population. Both theory and empirical evidence suggest that this impact varies with the initial extent to which liquidity constraints are binding. Suppose a market exists where VA-style mortgage innovations are superfluous because liquidity constraints are (just barely) not initially binding. Theoretically, we would expect to observe no impact on the ratio of values to rents.[28] The results suggest that in a

[28] In Table 5, we observe that increases in eligible veterans' share in low-rent markets have the effect of lowering housing values. This effect is present both in 1940 and 1970, suggesting that this pattern has more to do with nonrandom sorting of veterans than the VA mortgage program itself. The thought experiment conducted here starts with the presumption that the VA mortgage program would have no impact on prices in a housing market where liquidity constraints were not binding initially.

market with rent levels 10% higher than this unaffected city, extending VA-style mortgage benefits to an additional 10% of the population would increase the ratio of housing values to rents by about 6%.

Over the past twenty years, further innovations in the mortgage market have afforded a very large number of households access to borrowing regimes equivalent to, or in many cases more advantageous than, the VA mortgage program. Extrapolating from these results, extending VA borrowing privileges to the 60% of the population that was not eligible (as of 1970) would lead to value increases of 36% in metropolitan areas where rents exceeded those in a marginally unaffected city by 10%. Effects on housing values in low-rent areas, of course, would be decidedly more muted. Table 1 shows that relative to median rents, median housing values have increased by roughly 10 to 20 percent since 1970. While the evidence presented here is insufficient to directly link recent price increases to mortgage market innovations, the magnitude of the implied effect is consistent with such a link.

*C. Assessing the validity of veteran status as a proxy for access to credit*

Table 4 indicated that the impact of veteran status on home ownership was most prominent in 1970, the year examined in the preceding subsection. In other years, when veteran status conferred little if any true advantage in access to credit, the results in Table 5 should not be replicable. A strong resemblance between regressions using data from other years and those reported in Table 5 would support the hypothesis that veteran status captures some other characteristic of a housing market, rather than the density of householders with easy access to credit.

Table 6 displays coefficients drawn from four regression models that mirror the last reported

specification in Table 5, using data from the 1940, 1980, 1990 and 2000 IPUMS samples.[29] In each of the four data samples, higher veteran densities are associated with lower overall housing values.[30] Moreover, none of the four specifications feature a positive interaction term between median rent and veteran density. This may be somewhat surprising in the 1980 sample, since Table 4 indicated that VA mortgage eligibility conferred some advantage in that year. There are two related explanations for this finding. First, the overall density of veterans in the population declined from roughly 40% to 30% of all householders between 1970 and 1980, which could impact estimates if the true underlying effects are nonlinear. Second, veterans as a group were significantly older in 1980 than in 1970. In 1970, the mean age for veteran householders was 3.6 years lower than for non-veterans. By 1980, this pattern had reversed: the mean age of veteran householders was 3.1 years older than non-veterans. The density of veterans in age cohorts most traditionally associated with transitions to home ownership thus declined even more rapidly.

By 2000, the interaction between veteran density and median rent is negative and significant, indicating that housing now sells at a discount relative to rent levels in markets with a higher share of veterans. There are two possible explanations for this finding. First, the transition from a drafted to volunteer armed forces may have changed the relationship between veteran status and socioeconomic status. Veteran density may therefore mark socioeconomically disadvantaged markets in later data. Second, the comparative rarity of veterans in younger age cohorts implies that markets with higher veteran density may be declining areas where few young families are moving

[29] The 1940 specification more closely resembles the second regression model in Table 5, since structural characteristic variables are not available in that sample.

[30] Point estimates in the 1940 specification indicate that the predicted effect of an increase in veteran density is always negative, since the minimum value for ln(median rent) is roughly 2.

in.

If veteran density marks poor or declining metropolitan areas in 2000, might the same variable mark prosperous, growing areas in 1970? While the regressions in Table 5 control for a measure of recent growth, the concern that veteran status correlates with unobserved prosperity is serious enough to warrant additional attention. Table 7 investigates this concern with three additional specifications estimated using 1970 IPUMS data. The first tests to see whether faster-growing markets, as measured by the fraction of housing units built within the past ten years, have higher ratios of owner-occupied housing values to rents. The results indicate that this is not the case. The interaction between the recent growth variable and median rent is negative and statistically insignificant. Including this interaction term has essentially no impact on the significant, positive interaction between veteran density and median rent. The second and third specifications divide the sample evenly into markets that exhibited greater and lesser amounts of growth in the housing stock between 1960 and 1970. Within both subsamples, the significant interaction between veteran density and median rent persists. In fact, the point estimate is larger and of greater statistical significance within the subset of slower-growing metropolitan areas. Thus, the tendency for housing values to increase more rapidly with rent increases in markets with a high veteran tendency does not appear to be attributable to differential growth patterns across metropolitan areas.

**VI. Does easy credit harm the poor?**

The relaxation of liquidity constraints, by raising prices for durable goods, has the potential to harm some consumers. In the case of the mortgage market, where most innovations have involved

reducing the amount of savings necessary to purchase a home, the "harmed" consumers are those who had sufficient savings to participate in the mortgage market prior to the relaxation. Are these consumers drawn disproportionately from the upper or lower tails of the socioeconomic distribution? Table 8 analyzes this question empirically, analyzing cross-sectional variation in home ownership as of 1970. These probit specifications incorporate the same set of householder characteristics utilized in Table 4, adding the set of market characteristics used in Table 5. One additional market characteristic, the standard deviation of the log income distribution in each market, is added to capture the possibility that income inequality affects home ownership probabilities. Two separate variables are used to identify individuals of low socioeconomic status (SES). The first indicator is based on householders' reported occupation, and flags those working in jobs that are not ... The second is based on householder education, and flags those with no more than a high school diploma. Variation across markets in the extent of consumer access to credit is measured by the share of householders who are veterans eligible for the VA mortgage program.

The coefficient of greatest interest for this exercise is on the interaction between the low SES indicator and veteran density. In the first specification, which uses the occupational proxy for low SES, the interaction term is significant and negative, indicating that the gap in home ownership probabilities between householders of high and low SES is widest in markets with a high density of veterans. Note that the probability of home ownership is still higher for all householders in markets with high veteran density.[31]

The second regression model adds additional interaction terms, interacting the low SES

[31] This specification controls for veteran status as a householder characteristic, thus this result is independent of the fact that veterans themselves have higher home ownership probabilities.

indicator with the two other market characteristics showing a significant relationship with the probability of home ownership in the first regression: the shares of foreign born and college educated householders in the metropolitan population. The interaction between college graduate share and low SES proves to be statistically significant, indicating that the socioeconomic home ownership gap is largest in markets with a highly educated population. This term reduces the magnitude and significance of the veteran share-low SES interaction.

Results are generally quite similar when the occupational SES proxy is replaced with one based on education. The final pair of specifications in Table 8 show that adding interactions between low SES and college graduate share actually increases the magnitude of the the low SES-veteran share interaction term, while having no effect on the statistical significance. In both specifications, the veteran share-low SES interaction is negative and significant at the 10% level.

Overall, this analysis provides some suggestive evidence that the net impact of relaxing credit constraints can be regressive. Of course, the impact of the VA mortgage program on low-SES home ownership rates might be very different from the impact of a program targeted directly at lower income households. Given that very few government- or private-sector mortgage programs have any explicit means-tested component, and that such components usually consider measures of present rather than lifetime wealth, the possibility that recent innovations have hindered rather than promoted low income home ownership is a genuine concern worth further investigation (Goodman and Nichols 1997).

## VII. Conclusion

Previous research has frequently implicated borrowing constraints as significant obstacles

to home ownership. This literature, as well as the general literature on liquidity constraints and consumption, has to a surprising extent ignored the possibility that the relaxation of impediments to borrowing might have an inflationary effect on prices. Some authors have argued that elastic housing supply would dampen any price effects. This paper has shown, in a simple intertemporal model, that such an effect can indeed exist, regardless of the overall price elasticity of supply. Moreover, it has provided empirical evidence, centered around one of the great liquidity constraint-removing policy programs of the mid-twentieth century, that supports the existence of such an effect.

The implications of these price effects are quite noteworthy. The relaxation of borrowing constraints, by increasing the equilibrium price for durable goods such as housing, reduces the welfare of some households: those that were net purchasers of housing in the pre-relaxation equilibrium. In more familiar terms, the type of household that suffers from the relaxation of borrowing constraints is the type which has already saved for a down payment, only to witness the elimination of down payment requirements – and a concomitant increase in the price of owner occupied housing – prior to their purchase.

These results can be interpreted as a caveat to those interested in the policy goal of increasing home ownership rates, or rates of ownership of other durable goods. The relaxation of borrowing constraints, frequently touted as a central policy tool for achieving this goal, comes with a cost, even in the absence of borrower default risk. As theory and the evidence drawn from the VA mortgage program indicates, easier credit for some can make prices higher for all.

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Table 1: Housing prices and home ownership rates, 1940-2000

| Year | Median value | Median monthly contract rent | Ratio of median value to median monthly contract rent | National home ownership rate |
|---|---|---|---|---|
| 1940 | $2,750 | $20 | 137.5 | 43.6% |
| 1950 | — | — | — | 55.0 |
| 1960 | 11,250 | 55 | 182-227 | 61.9 |
| 1970 | 16,250 | 87 | 172-201 | 62.9 |
| 1980 | 47,500 | 195 | 231-256 | 64.4 |
| 1990 | 85,000 | 362 | 221-249 | 64.2 |
| 2000 | 112,500 | 500 | 200-250 | 66.2 |

Note: Median value sample consists of owner-occupied nonfarm single family units, excluding vacant-for-sale and condominium properties, and properties with commercial uses. In 1940, values are for all nonfarm owner-occupied units; owners of multifamily or commercial-use units were instructed to report only the market value of their own unit. Median rent sample uses contract rent observations for all nonfarm rental units with cash rent. Ranges reported for ratio of median value to median rent when value data are reported in intervals only.
Source: IPUMS samples, 1940-2000. The 1950 IPUMS sample contains no information on housing values or rents.

Table 2: Important events in the history of the mortgage market

| | |
|---|---|
| 1920s | Typical mortgage lasts 5-10 years, requires at least 30% down payment, and is less than fully amortized. |
| 1932 | 250,000 nonfarm foreclosures nationwide. |
| 1933 | Home Owners Loan Corporation established, makes fully amortized 20-year loans. |
| 1934 | FHA establishes program of insuring against homeowner default on long-term, fixed-rate loans. Enables down payments of less than 10% of property value, 25-30 year loan periods with full amortization. Borrowers pay a mortgage insurance premium in exchange for coverage. |
| 1938 | Fannie Mae established, creates secondary market for mortgage loans that "conform" to underwriting standards covering borrower creditworthiness and property characteristics. |
| 1944 | Servicemen's Readjustment Act creates Veterans Administration mortgage insurance program, which enables qualifying veterans to purchase a home with no down payment and no mortgage insurance premiums. |
| 1968 | Fannie Mae privatized, Ginnie Mae established to fund FHA and VA mortgage programs. |
| 1970 | Freddie Mac established as an additional secondary market conduit. |
| 1982 | Beginning of rapid growth in the secondary market for "non-conforming" loans; mortgages that Fannie Mae and other agencies are unwilling to securitize. |
| 1992 | Federal Housing Enterprises Financial Safety and Soundness Act directs Fannie Mae and Freddie Mac to acquire more loans made to low-income borrowers. This directive leads to more flexible terms for "conforming" loans. |
| 1993 | $100 billion worth of non-conforming loans securitized in the secondary market. |
| 1994 | Establishment of loan programs that allow qualified buyers to borrow up to 125% of property value. |

Sources: Jackson (1985), Bruskin et al. (2000), Listokin et al. (2001), Martinez (2000).

Table 3: Relationship between median rents and owner-occupied house values, 1940-2000

| | Dependent variable: ln(value) for owner-occupied units in year: | | | | | | | |
|---|---|---|---|---|---|---|---|---|
| Independent variable | 1940 | 1940 | 1970 | 1970 | 1970 | 1980 | 1990 | 2000 |
| ln(median rent) | 0.915<br>(0.074) | 0.297<br>(0.101) | 1.068<br>(0.146) | 1.174<br>(0.157) | 1.103<br>(0.159) | 0.921<br>(0.130) | 1.948<br>(0.134) | 1.815<br>(0.142) |
| Intercept | 5.173<br>(0.248) | — | 5.055<br>(0.699) | — | — | — | — | — |
| Housing market characteristic controls? | No | Yes | No | Yes | Yes | Yes | Yes | Yes |
| Median rent adjusted for structural characteristics? | No | No | No | No | Yes | Yes | Yes | Yes |
| Structural characterstic controls? | No | No | No | No | Yes | Yes | Yes | Yes |
| N | 66,050 | 66,050 | 192,581 | 192,581 | 192,581 | 291,624 | 385,877 | 309,363 |
| Number of metropolitan areas | 137 | 137 | 124 | 124 | 124 | 286 | 297 | 106 |
| $R^2$ | 0.040 | .045 | 0.143 | 0.153 | 0.533 | 0.520 | 0.615 | 0.517 |

Note: Huber-White robust standard errors in parentheses. Regression observations are weighted by the size of the sample of rental units used to create the median rent measure. Sample consists of owner-occupied nonfarm housing units. In 1940, sample is restricted to units where the householder is a sample line person; owners of multifamily dwellings were instructed to report only the value of their own dwelling. In later years, multifamily dwellings are excluded from the sample. Housing market characteristic controls include median householder age, percent of householders who are college-educated, percent of householders in the labor force that work in manufacturing industries, median family income, region effects, and in 1970-2000, percent of housing units built in the past ten years. Structural characteristic controls include categorical variables for number of rooms, year structure built, presence of basement, number of units in structure (rental units only), whether unit is detached (owner-occupied units only), number of units at address (rental units only), number of bathrooms, and presence of window or central air conditioning. All coefficients reported in this table are significantly different from zero at the 1% level.

Table 4: Did the VA mortgage program give veterans an advantage?

| Independent Variable | Probit regression results: Dependent variable equal to one if respondent owns home | | | | |
|---|---|---|---|---|---|
| | 1940 | 1970 | 1980 | 1990 | 2000 |
| Age | 0.025 | 0.039 | 0.034 | 0.037 | 0.034 |
| | (0.001) | ($4.90*10^{-4}$) | (2.77*10-4) | (2.64*10-4) | (2.76*10-4) |
| Age squared | $-1.50*10^{-4}$ | $-2.95*10^{-4}$ | -2.42*10-4 | -2.57*10-4 | -2.35*10-4 |
| | ($1.33*10^{-5}$) | ($5.44*10^{-6}$) | (2.75*10-6) | (2.49*10-6) | (2.56*10-6) |
| Married | 0.065 | 0.284 | 0.253 | 0.211 | 0.201 |
| | (0.007) | (0.004) | (0.002) | (0.002) | (0.002) |
| Female | -0.055 | 0.092 | 0.026 | -0.009 | -0.006 |
| | (0.009) | (0.004) | (0.002) | (0.002) | (0.002) |
| Black | -0.125 | -0.216 | -0.194 | -0.180 | -0.174 |
| | (0.007) | (0.003) | (0.002) | (0.002) | (0.002) |
| Foreign Born | -0.029 | -0.192 | -0.212 | -0.213 | -0.229 |
| | (0.006) | (0.004) | (0.003) | (0.002) | (0.002) |
| ln(family income) | 0.067 | 0.092 | 0.131 | 0.138 | 0.139 |
| | (0.003) | (0.001) | (0.001) | (0.001) | (0.001) |
| Household size | 0.028 | 0.059 | 0.039 | 0.018 | 0.025 |
| | (0.002) | (0.001) | (0.001) | (0.001) | (0.001) |
| Northeast region | -0.127 | -0.092 | -0.089 | -0.018 | -0.036 |
| | (0.007) | (0.003) | (0.002) | (0.002) | (0.002) |
| South region | -0.014 | 0.040 | 0.057 | 0.073 | 0.091 |
| | (0.008) | (0.003) | (0.002) | (0.002) | (0.002) |
| Midwest region | -0.040 | 0.041 | 0.055 | 0.080 | 0.099 |
| | (0.007) | (0.003) | (0.002) | (0.002) | (0.002) |
| Veteran | -0.012 | -0.015 | -0.020 | -0.027 | -0.018 |
| | (0.007) | (0.010) | (0.009) | (0.013) | (0.017) |
| Veteran eligible for VA mortgage | — | 0.076 | 0.042 | 0.011 | 0.011 |
| | | (0.010) | (0.009) | (0.013) | (0.017) |
| N | 44,449 | 279,505 | 549,888 | 599,863 | 491,500 |
| Pseudo-$R^2$ | 0.117 | 0.217 | 0.252 | 0.244 | 0.253 |

Note: Table entries represent the marginal change in predicted probability associated with a unit change in the independent variable when all other covariates are set equal to their means. Veterans designated eligible for VA mortgages in 1970 are those reporting service in World War II, Korea or Vietnam, or those under 50 reporting peacetime military service only. Sample consists of metropolitan area-resident nonfarm householders in the merged 1940 (sample line persons only), 1970 (Form 2 metro sample), 1980, 1990 and 2000 IPUMS. All coefficients in this table except South region in 1940 and Veteran in the first, second, and fifth column are significantly different from zero at the 0.1% level.

Table 5: Housing values and the density of veterans in the population, 1970

| Independent variable | Dependent variable: ln(housing value) | | | |
|---|---|---|---|---|
| ln(median rent) | 0.878*** | -1.978** | 0.759*** | -1.693*** |
| | (0.143) | (0.733) | (0.113) | (0.493) |
| Share of veterans in the population | 2.279*** | -35.58*** | -2.661*** | -2.895*** |
| | (0.567) | (9.123) | (0.512) | (0.401) |
| Veterans' share*ln(median rent) | — | 7.206*** | — | 6.242*** |
| | | (1.975) | | (1.292) |
| Median householder age | -0.004 | -0.003 | -0.007 | -0.004 |
| | (0.008) | (0.010) | (0.007) | (0.007) |
| Share of householders with college degree | 2.046** | 2.671*** | 0.915 | 1.548** |
| | (0.858) | (0.809) | (0.690) | (0.671) |
| Share of housing units built in past 10 years | -0.521 | -0.782** | -0.764** | -0.968*** |
| | (0.356) | (0.346) | (0.317) | (0.283) |
| Share of householders employed in manufacturing industries | 0.294 | 0.610** | 0.209 | 0.546** |
| | (0.372) | (0.294) | (0.295) | (0.218) |
| Log of median family income | 0.166 | 0.010 | 0.230 | 0.041 |
| | (0.284) | (0.282) | (0.225) | (0.225) |
| Northeast region | -0.078 | -0.159*** | -0.132*** | -0.209*** |
| | (0.062) | (0.060) | (0.039) | (0.043) |
| Midwest region | -0.052 | -0.066** | -0.153*** | -0.165*** |
| | (0.036) | (0.031) | (0.026) | (0.023) |
| South region | -0.158*** | -0.142*** | -0.260*** | -0.244*** |
| | (0.056) | (0.054) | (0.046) | (0.044) |
| Median rent adjusted for structural characteristics? | No | No | Yes | Yes |
| Structural characteristic controls? | No | No | Yes | Yes |
| N | 192,581 | 192,581 | 192,581 | 192,581 |
| $R^2$ | 0.160 | 0.166 | 0.543 | 0.548 |

Note: Huber-White robust standard errors in parentheses. Regression observations are weighted by the size of the sample of rental units used to create the median rent measure. Sample consists of owner-occupied nonfarm housing units. Multifamily dwellings are excluded from the sample. Housing market characteristic controls include median householder age, percent of householders who are college-educated, percent of householders in the labor force that work in manufacturing industries, region effects, and percent of housing units built in the past ten years. Structural characteristic controls include categorical variables for number of rooms, year structure built, presence of basement, whether unit is detached, number of bathrooms, and presence of window or central air conditioning.

*** denotes a coefficient significant at the 1% level, ** the 5% level, * the 10% level.

Table 6: Housing values and the density of veterans in the population in other years

| Independent variable | 1940 | 1980 | 1990 | 2000 |
|---|---|---|---|---|
| ln(median rent) | 0.636*** | 1.266 | 1.687*** | 2.603*** |
| | (0.187) | (0.859) | (0.113) | (0.469) |
| Share of veterans in the population | 6.234*** | -2.371*** | -2.627*** | -1.581*** |
| | (6.633) | (0.748) | (0.713) | (0.528) |
| Veterans' share*ln(median rent) | -3.518 | -2.031 | -0.731 | -5.230*** |
| | (2.261) | (2.750) | (2.239) | (2.050) |
| Median householder age | -0.012 | 0.009 | 0.011 | 0.025** |
| | (0.008) | (0.006) | (0.010) | (0.010) |
| Share of householders with college degree | 4.860*** | 1.480*** | 0.252 | 1.154*** |
| | (1.314) | (0.442) | (0.487) | (0.300) |
| Share of housing units built in past 10 years | — | -1.033*** | -0.973*** | -1.178*** |
| | | (0.230) | (0.285) | (0.271) |
| Share of householders employed in manufacturing industries | -0.606 | -0.325 | -0.584 | 0.196 |
| | (0.196) | (0.270) | (0.368) | (0.354) |
| Log of median family income | 0.496 | -0.012 | -0.500** | -0.748*** |
| | (0.109) | (0.263) | (0.236) | (0.156) |
| Northeast region | 0.026 | -0.480*** | -0.074 | -0.250*** |
| | (0.046) | (0.442) | (0.065) | (0.034) |
| Midwest region | -0.005 | -0.232*** | -0.246*** | -0.125*** |
| | (0.036) | (0.056) | (0.061) | (0.045) |
| South region | -0.131* | -0.397*** | -0.251*** | -0.301*** |
| | (0.068) | (0.048) | (0.050) | (0.035) |
| Median rent adjusted for structural characteristics? | No | Yes | Yes | Yes |
| Structural characteristic controls? | No | Yes | Yes | Yes |
| N | 66,050 | 291,624 | 385,877 | 309,363 |
| $R^2$ | 0.046 | 0.525 | 0.621 | 0.520 |

Note: Huber-White robust standard errors in parentheses. Regression observations are weighted by the size of the sample of rental units used to create the median rent measure. Sample consists of owner-occupied nonfarm housing units. In 1940, sample is restricted to units where the householder is a sample line person; owners of multifamily dwellings were instructed to report only the value of their own dwelling. In later years, multifamily dwellings are excluded from the sample. Housing market characteristic controls include median householder age, percent of householders who are college-educated, percent of householders in the labor force that work in manufacturing industries, region effects, and after 1940, percent of housing units built in the past ten years. Structural characteristic controls include categorical variables for number of rooms, year structure built, presence of basement, whether unit is detached, number of bathrooms, and presence of window or central air conditioning.

*** denotes a coefficient significant at the 1% level, ** the 5% level, * the 10% level.

Table 7: Is the veteran effect a boomtown phenomenon?

| Independent variable | Complete sample | High-growth cities | Low-growth cities |
|---|---|---|---|
| ln(median rent) | -1.765** | -2.162 | -3.162*** |
| | (0.865) | (1.326) | (0.636) |
| Share of veterans in the population | -2.780*** | -2.580*** | -2.208*** |
| | (0.469) | (0.912) | (0.388) |
| Veterans' share*ln(median rent) | 6.943*** | 6.953* | 10.29*** |
| | (2.190) | (3.556) | (1.709) |
| Share of housing units built in past 10 years | -0.882 | -0.397 | -3.084*** |
| | (0.424) | (0.563) | (0.871) |
| Share of housing units built in past 10 years*ln(median rent) | -0.771 | — | — |
| | (0.847) | | |
| Additional housing market characteristics? | Yes | Yes | Yes |
| Median rent adjusted for structural characteristics? | Yes | Yes | Yes |
| Structural characteristic controls? | Yes | Yes | Yes |
| N | 192,581 | 76,295 | 116,286 |
| $R^2$ | 0.532 | 0.596 | 0.523 |

Note: Huber-White robust standard errors in parentheses. Regression observations are weighted by the size of the sample of rental units used to create the median rent measure. Sample consists of owner-occupied nonfarm single-family housing units in the 1970 IPUMS. Additional housing market characteristic controls include median householder age, percent of householders who are college-educated, percent of householders in the labor force that work in manufacturing industries, log of median family income, and region effects. Structural characteristic controls include categorical variables for number of rooms, year structure built, presence of basement, whether unit is detached, number of bathrooms, and presence of window or central air conditioning. High- and low-growth cities are consist of the 62 metropolitan areas with above- and below-average shares of housing built in the past ten years. The cutoff point between categories is 27.35%.
*** denotes a coefficient significant at the 1% level, ** the 5% level, * the 10% level.

Table 8: The distributional impact of selectively relaxing credit constraints

| | Dependent Variable: Home ownership indicator | | | |
|---|---|---|---|---|
| Independent Variable | SES Proxy: Occupation | | SES Proxy: Education | |
| Low SES indicator | 0.081 | 0.120** | 0.060 | 0.155** |
| | (0.036) | (0.051) | (0.039) | (0.061) |
| Market characteristics | | | | |
| VA mortgage-eligible veterans as share of householders | 1.63** | 1.57** | 1.56** | 1.62** |
| | (0.382) | (0.366) | (0.377) | (0.369) |
| Median age | 0.005 | 0.005 | 0.005 | 0.005 |
| | (0.011) | (0.011) | (0.011) | (0.011) |
| Foreign-born share of householders | -0.902** | -0.890** | -0.905** | -0.825** |
| | (0.296) | (0.283) | (0.295) | (0.287) |
| ln(median family income) | 0.250 | 0.250 | 0.253 | 0.253 |
| | (0.241) | (0.240) | (0.241) | (0.240) |
| standard deviation of ln(family income) | 0.163 | 0.185 | 0.171 | 0.188 |
| | (0.320) | (0.323) | (0.320) | (0.324) |
| Percent of housing units built in last 10 years | 0.271 | 0.275 | 0.271 | 0.272 |
| | (0.238) | (0.239) | (0.239) | (0.240) |
| College graduates as share of householders | -1.325** | -0.975** | -1.32** | -1.10** |
| | (0.513) | (0.454) | (0.513) | (0.446) |
| Percent of householders employed in manufacturing industries | 0.157 | 0.157 | 0.157 | 0.157 |
| | (0.158) | (0.158) | (0.158) | (0.158) |
| Interaction of Low SES indicator with | | | | |
| VA mortgage-eligible veterans as share of householders | -0.292** | -0.178 | -0.197* | -0.264* |
| | (0.101) | (0.132) | (0.101) | (0.160) |
| Foreign-born share of householders | — | -0.001 | — | -0.100 |
| | | (0.073) | | (0.088) |
| College graduates as share of householders | — | -0.975** | — | -0.330* |
| | | (0.454) | | (0.179) |
| Additional householder characteristics | Yes | Yes | Yes | Yes |
| Region effects | Yes | Yes | Yes | Yes |
| N | 337,675 | 337,675 | 337,675 | 337,675 |
| Pseudo-$R^2$ | 0.253 | 0.253 | 0.252 | 0.252 |

Note: Table entries reflect probit coefficients scaled to indicate marginal impact on predicted probability when other covariates are set equal to their respective means. Standard errors, corrected for grouped observations, in parentheses. Observations are weighted by the number of observations used to calculate the market characteristic variables. Additional householder characteristics include age, age squared, whether married, whether female, whether black, whether foreign born, logarithm of household income, household size, and whether a VA mortgage-eligible veteran. Data source: IPUMS.

** denotes a coefficient significant at the 1% level, * the 5% level.

**Figure 1: The price of owner-occupied housing relative to all personal consumption**

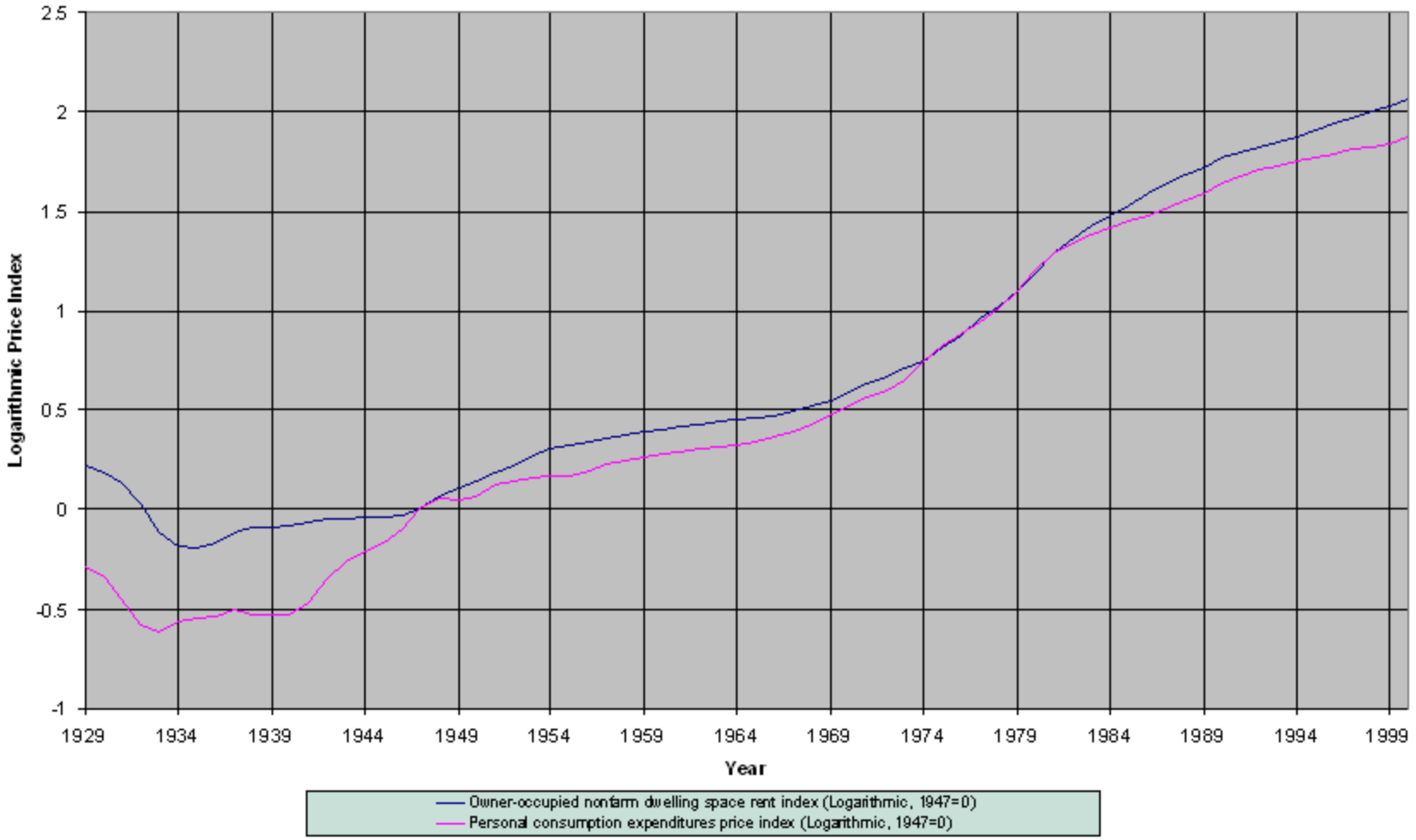

Table A1: Summary statistics for regression covariates

| | 1940 | | | | | 1970 | | | | |
|---|---|---|---|---|---|---|---|---|---|---|
| Variable | N | mean | std. dev. | minimum | maximum | N | mean | std. dev. | minimum | maximum |
| Householder owns home | 44,449 | 0.303 | — | — | — | 279,505 | 0.569 | — | — | — |
| Age | 44,449 | 42.701 | 12.468 | 15 | 100 | 279,505 | 42.629 | 13.636 | 14 | 100 |
| Married | 44,449 | 0.779 | — | — | — | 279,505 | 0.756 | — | — | — |
| Female | 44,449 | 0.139 | — | — | — | 279,505 | 0.160 | — | — | — |
| Black | 44,449 | 0.081 | — | — | — | 279,505 | 0.107 | — | — | — |
| Foreign Born | 44,449 | 0.198 | — | — | — | 279,505 | 0.085 | — | — | — |
| ln(wage and salary income) | 44,449 | 6.991 | 0.849 | 0.000 | 8.517 | 279,505 | 8.809 | 0.873 | 3.912 | 10.820 |
| Household size | 44,449 | 2.820 | 1.547 | 1 | 29 | 279,505 | 3.316 | 1.785 | 1 | 26 |
| Veteran | 44,449 | 0.091 | — | — | — | 279,505 | 0.463 | — | — | — |
| Veteran eligible for VA mortgage | 44,449 | 0.000 | — | — | — | 279,505 | 0.452 | — | — | — |
| ln(owner-occupied housing unit value) | 66,050 | 8.046 | 0.918 | 0.000 | 11.362 | 192,581 | 9.844 | 0.584 | 7.824 | 10.820 |
| ln(median rent) | 137 | 2.923 | 0.315 | 2.079 | 3.689 | 124 | 4.446 | 0.235 | 3.912 | 4.942 |
| Veterans' share | 137 | 0.080 | 0.026 | 0.017 | 0.170 | 124 | 0.416 | 0.033 | 0.283 | 0.508 |