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The Bank's Market Power and the Interest-Rate Elasticity of Demand for Housing: An Econometric Study of Discrimination on French Mortgage Data

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ABSTRACT. – We specify several variants of a structural econometric model explaining mortgage interest rates and loan sizes simultaneously. The models are estimated by simultaneous equation methods with a sample of loan files originated from a French mortgage lender. They yield estimates of the interest-rate elasticity of the demand for housing. Different occupational status groups happen to have different values of structural preference parameters. We show how these differences translate into differential treatment of socio-economic groups by the banker.

Le pouvoir de marché des banques et l'élasticité aux taux d'intérêt de la demande de logement

RÉSUMÉ. – Nous spécifions plusieurs variantes d'un modèle économétrique structurel expliquant les taux d'intérêt et les montants de crédit simultanément. Les modèles sont estimés par des méthodes adaptées aux équations simultanées avec un échantillon de dossiers de crédit émanant d'un réseau de prêteurs hypothécaires français. Ils fournissent des estimations de l'élasticité par rapport au taux d'intérêt de la demande de logement. Ces élasticités varient avec la catégorie socioprofessionnelle de l'emprunteur : des groupes différents ont des paramètres structurels de préférence différents. Nous montrons comment ces différences se traduisent par un traitement différencié des clients de la part du banquier.

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1 Introduction

The basic goal of this paper is to describe the pricing of mortgage credit in France, using a price discrimination model where lenders charge different interest rates to different social groups. The underlying assumption of the empirical work is that the interest-rate elasticity of demand for housing varies with the occupational status of borrowers. The data used for estimation is a sample of clients of a French network of mortgage lenders. To the best of our knowledge, interest rate discrimination in the mortgage lending sector has not been studied with the help of structural econometric methods and samples of individual loans.

We specify and estimate variants of a structural model, explaining variations in the amount lent and the interest rate charged to borrowers simultaneously. Assuming that observed households are liquidity-constrained and cannot borrow against future income, we use the constant repayment (or self-amortizing annuity) of the (fixed rate) mortgage as a measure of loan "price".

In a first version of the model, based on the classic idea of third-degree price discrimination, the loan price, as defined above, determines the borrower's demand for a loan, along with income, personal and socioeconomic characteristics. Assuming that the lender is endowed with a form of market power, the equilibrium loan price is in turn derived from profit maximization: it is a function of borrower characteristics. The markup added to the banker's cost of funds depends on a risk premium and on the borrower's price-elasticity of demand. We obtain a loglinear, simultaneous equation model in structural form, which is estimated by means of the classic 2SLS and 3SLS methods. The model is also estimated separately on subsamples of borrowers belonging to four categories: blue-collar workers, white-collar workers, intermediate professions, and executives. We find a ranking of elasticities, executives exhibiting a higher price-elasticity of demand than workers. This leads the banker to optimally discriminate in a striking way: the model predicts a much higher margin for workers than for executives.

A second version of the model is based on the idea of "competition in mortgage loan contracts" by lenders, based on observable borrower characteristics.[1] Two equations describe the credit market equilibrium. Mortgage pricing is now given by a zero-profit condition applied to each observable category of borrowers: the interest rates charged are expressed as the sum of a risk premium and a cost of funds. The loan amount is determined by a contract optimality condition: in equilibrium, no lender can offer a more advantageous price-quantity (*i.e.*, loan amount and interest rate) pair to any type of borrower without making losses in expectation. These assumptions lead to a bivariate, nonlinear model, which is estimated with the help of the standard full information maximum likelihood method. Estimation results confirm the fact that structural preference parameters seem to vary from one

1. Our model thus represents third-degree price discrimination rather than second-degree price discrimination (or screening under incomplete information à la Rothschild-Stiglitz).

occupational group to the other, and offer a basis for interest-rate and loan size discrimination.

In the final section, we discuss possible relationships of our study with the literature on (racial) discrimination in mortgage lending in the United States. Our models could be adapted to exploit other data sets, and used to test for the presence of racial discrimination.[2] The theory of discrimination in mortgage markets has recently been studied by (among other studies), CHARI and JAGANNATHAN [1989], BRUECKNER [1994], CALEM and STUTZER [1995], STANTON and WALLACE [1997], and BRUECKER [2000]. The bulk of the literature on discrimination in mortgage markets is empirical (see references in Section 5 below).

In the following, Section 2 is devoted to a presentation of the data; Section 3 presents the monopolistic model and its estimations. The analysis and estimation of the nonlinear market equilibrium model is presented in Section 4. Relationships with the literature on racial discrimination are discussed in Section 5.

2 The Data

We have used a sample of observations on the clients of a French mortgage lender, the Crédit Hypothécaire de France (a nickname), hereafter CHF. The CHF is in fact a network of building societies, scattered on the French territory, the BSs. These local BSs have independent application screening and interest rate policies; they own in common a financial institution, which borrows money on national and international bond markets, and provides funds to the BSs. The CHF is a prudent and profitable institution, with a long history and a solid reputation. The BSs do not securitize their loans. Rating agencies have granted a very high note (AA+) to the CHF, so that the institution's cost of funds is well approximated by, and closely parallels, the long-term rate on French state bonds (the “OAT” rate), with an almost constant difference of a few base points. Although the CHF has a special legal status, it is fair to describe the behavior of the local BSs as profit maximization. Until 1995, when the French government reformed its housing policy, the CHF had the privilege of distributing a particular kind of state subsidized home loans. This privilege has disappeared today, since all commercial banks can now initiate the same subsidized loans, but the CHF network has developed a strong expertise in mortgage lending to the working class, and a goodwill in accordance with this specialization. Its clientele is composed of a vast majority of modest income employees and workers. It is likely that many of the CHF clients would see their applications rejected elsewhere.

On top of distributing state subsidized loans, the characteristics of which are tightly regulated, the CHF also supplies the so-called “free loans”, which are unregulated, ordinary mortgages. Until recently, the vast majority of these

2. In our data set, race is not observed.

mortgages have been classic, fixed rate, fixed repayment mortgages. The French mortgage law is in a sense simpler than the U.S. legal environment, since the borrowers' liability is *not* limited to the value of the house (lenders can pursue other borrower assets to mitigate default-related losses). In addition, house prices have not decreased very much in the provincial regions, which are the geographical origin of the sampled borrowers, during the observation period. It follows that strategic default (or the exercise of the default option) is not empirically relevant in the sample. In practice, mortgage defaults seem to be triggered by consumer insolvency, mostly due to loss of income. A form of unemployment insurance of mortgage loans does indeed exist, but it is not compulsory, it is expensive, and limited in scope. These loans can in principle be renegotiated, the prepayment penalty being in all BSs around 3% of the principal's remaining value. For the econometric investigations below, we have used a sample of 2610 observations on *accepted* free loans, originated from various BSs across France between 1989 and 1994. We have eliminated the subsidized loans. There is no information on rejected applications, and no observations of default or of repayment "incidents". Each observation corresponds to a file, including, 1°) the amount of the loan, 2°) the loan interest rate (including insurance), 3°) the downpayment (savings used to buy the house by the borrower), 4°) the starting year, 5°) the loan term, 6°) the borrower's yearly wage, 7°) the age of the borrower, 8°) the family size, 9°) the borrower's occupational status, falling into 6 categories, and 10°) the geographical location of the BS granting the loan. We kept only four of the occupational status categories, the blue-collar workers (1180 observations), the white-collar employees (908 observations), the so-called intermediate professions (363 observations), and the executives (159 observations). The models presented below have been constructed to be estimated with this limited set of information. They can of course easily be adapted to use more explanatory variables. Table 1 provides descriptive statistics on the data. The amounts lent are not very high. The downpayment ratios, that is, (downpayment /loan + downpayment) are around 17%. The loan terms are distributed between 1 year and 20 years, with 85% of the loans having a term of 10, 15 or 20 years (and 42% of the loans having a term of 15 years). The nominal interest rates are very high with a mean value of 11,7%, but the real interest rates where also very high at the beginning of the nineties in France, the inflation rate being already quite low around 2% per year. The interesting aspect of the data is the substantial variance of the loan rates. This will allow the estimation of an interest-rate elasticity of the demand for housing, and of a risk-premium function, in spite of the fact that the observation period is short. Another striking fact is the markup on state bond rates, which is equal to 3 percentage points in the average. The lenders seem to exert a form of market power.[3]

3. According to the CHF management, the rate of problem loans is of the order of 1%. Mortgage foreclosure is a rare phenomenon. Even if this figure is underestimated, it does not seem to justify the observed markups on the long-term State bond ("OAT") rate.

TABLE 1
Descriptive Statistics

	4 Categories Pooled	Executives	Intermediate Professions	White Collars	Blue Collars
Number of Observations	2610	159	363	908	1180
Loan Size in euros – average – stand. dev.	 33540 18284	 44055 35705	 35188 23550	 35599 16779	 31569 12745
Downpayment Ratio – average – stand.dev.	 0.172 0.182	 0.297 0.271	 0.247 0.232	 0.162 0.174	 0.139 0.137
Loan Term in years – average – stand.dev.	 14.89 4.08	 13.08 4.05	 13.04 3.88	 14.91 4.02	 15.68 3.94
Loan Interest Rate – average – stand.dev.	 11.658 1.739	 10.316 2.076	 10.847 1.951	 11.791 1.608	 11.986 1.545
Cost of Funds – average – stand.dev.	 8.330 1.238	 7.888 1.038	 8.028 1.033	 8.257 1.244	 8.538 1.275
Yearly Wage in euros – average – stand.dev.	 20011 10936	 30671 20743	 22900 12976	 21018 10760	 16911 6160

3 The Monopolistic Model

Our first approach of the data will be to estimate a bivariate model explaining the interest rate and the amount lent simultaneously, as a function of borrower characteristics, and of the banker's cost of funds. We will study two variants of the model, hereafter called Model A and Model B. Both models reflect the view that the banker has exerted some market power while charging an interest rate, and that this practice is a form of third-degree price discrimination, made possible because the demand for housing's interest-rate elasticity varies between borrower groups. Another important assumption is that borrowers are liquidity-constrained (see DEATON [1992]).

3.1 Model A: Specification

The vast majority of our observations corresponding to workers or, at best, to middle class wage-earners, we assume that borrowers are severely liquidity-constrained: their savings will be used in full as a downpayment, denoted A, to purchase the house, of size H, with the help of a loan L. We have

$H = L + A$. These quantities are denominated in euros. The house is used as a collateral for the loan, and our workers cannot borrow against future wages. They therefore face a budget constraint at each period.[4] Their consumption is simply their wage W, minus mortgage payments, PL. A household who borrows for T years at rate r will repay each year, per euro borrowed, a self-amortizing annuity proportional to,

$$P = P(r,T) \equiv \frac{1 - \frac{1}{1+r}}{1 - \left(\frac{1}{1+r}\right)^T}. \tag{1}$$

This variable P is the only thing which matters for the liquidity-constrained borrower, and the interest rate r should not intervene directly, but only through P. The demand for loans is therefore assumed to depend on P. We assume that the demand for loans of a borrower with observable characteristics (x, W, A) writes,

$$L = \Phi(x) W^{\gamma} A^{\sigma} P^{-\varepsilon}, \tag{2}$$

where $\Phi(x)$ is a function of observable characteristics x; and γ, σ, and ε are positive parameters.

Assume that the banker has a very conservative interest-rate risk policy (which is indeed the case of the CHF), and matches the reimbursement flows on bonds, bearing interest rate i, with their clients' reimbursements. Then, the following variable C, defined as,

$$C = P(i,T), \tag{3}$$

is an appropriate measure of the lender's cost per euro lent, which is commensurate with P. To avoid confusion in the following, we rename the above function: $C(i,T) \equiv P(i,T)$.

Now to keep the model simple, we assume that the banker applies a discount factor π to repayments P, to take into account the risk of default on mortgage payments. Factor π can be interpreted as a "probability of repayment", and is assumed to depend on borrower characteristics as follows,

$$\pi = (1/\mu) F^{\beta} E^{-\lambda} C^{1-\delta}, \tag{4}$$

where $F = A/L$ is the downpayment ratio, and $E = L/W$ is the borrower's "effort ratio". Both variables are supposed to capture part of the risk: we assume that μ, β, δ and λ are positive parameters. Note that the usual initial loan-to-value ratio is $L/(L + A)$; the inverse LTV ratio is therefore equal to

4. We assume that all borrowers are in a corner solution, *i.e.*, are liquidity-constrained. This assumption might be too extreme for the high income categories. A model with two different regimes would be much more complicated. In addition, in our data set, the number of observations on high income borrowers would be too small to allow for a convincing estimation of the more complicated model.

$1 + F$. The classic effort ratio would be PL/W instead of L/W, but our simple formulation is enough to capture the idea that a heavy debt burden increases the risk of repayment problems. To increase our specification's flexibility, we add C as a variable: a rise in the cost of funds could increase risk, or it could also be the case that CL/W (instead of simply L/W) plays a role in explaining default risk.[5]

We now define the banker's profit on a borrower characterized by (x, W, A), denoted B, as follows,

$$B = (\pi P - C)L. \tag{5}$$

The bank is assumed to maximize B with respect to P, for each borrower type. We treat the loan term T as an exogenous variable here (it is then equivalent to a borrower characteristic).

To compute the first-order condition for profit maximization, note first that, combining the expressions of π and L, we get,

$$\pi = (1/\mu)A^{\beta}W^{\lambda}C^{1-\delta}L^{-(\beta+\lambda)} = ZP^{\varepsilon(\beta+\lambda)}, \tag{6}$$

where, $Z \equiv (1/\mu)\Phi^{-(\beta+\lambda)}A^{\beta-\sigma(\beta+\lambda)}C^{1-\delta}W^{\lambda-\gamma(\beta+\lambda)}$ is a function of exogenous variables. Using the notation $K \equiv \Phi W^{\gamma}A^{\sigma}$, we get

$$B = (ZP^{1+\,\varepsilon(\beta+\lambda)} - C)KP^{-\varepsilon}. \tag{7}$$

Maximization of this expression with respect to P yields, after simplifications, a modified form of the classic Lerner pricing formula,

$$\frac{P - \frac{C}{\pi}}{P} = \frac{1}{\varepsilon} + \beta + \lambda, \tag{8}$$

or, equivalently, $P = (C/\pi)(1 - (1/\varepsilon) - (\beta + \lambda))^{-1}$. Substituting the expression for π, this can finally be rewritten,

$$P = \Psi C^{\delta}\left(\frac{A}{L}\right)^{-\beta}\left(\frac{L}{W}\right)^{\lambda}, \tag{9}$$

where by definition,

$$\Psi = \frac{\mu}{1 - (1/\varepsilon) - (\beta + \lambda)}. \tag{10}$$

5. The inclusion of C in the π function could of course capture adverse selection phenomena à la STIGLITZ and WEISS, or C plays a role in the effort ratio. We should then test wether $\delta = 1$ or $\delta = 1 + \lambda$. None of these assumptions are confirmed by the data, and it follows that our best justification for introducing C in the π function is empirical.

Clearly, we must have $P > 0$, and we must therefore check that,

$$\varepsilon > \frac{1}{1-\beta-\lambda}. \tag{11}$$

We now have an equation giving P as a log-linear function of L and exogenous variables (x, W, A, C) and L is specified as a log-linear function of P and the same exogenous variables. Denote the logarithms of the variables as lower-case letters as follows, $\ell = \ln(L)$, $p = \ln(P)$, $a = \ln(A)$, etc. Taking logs and adding random disturbances, we get the following log-linear econometric specification,

$$\ell = \varphi(x) + \gamma w + \sigma a - \varepsilon p + u, \tag{12}$$

$$p = \psi + \delta c - \beta(a - \ell) + \lambda(\ell - w) + v, \tag{13}$$

where (u, v) are random perturbations, and where theory suggests that $e^{\psi} = \mu/(1 - 1/\varepsilon - \beta - \lambda)$. This constraint on parameters is simply a way of identifying μ. We of course compute $P = P(r, T)$ from the observations of (r, T) and $C = C(i, T)$ from the observations of (i, T). This linear system is a classic simultaneous equations system, which can be estimated by instrumental variable methods, to take care of endogeneity problems. To fully specify the model we have set,

$$\varphi(x) = \varphi_0 + \varphi_1 N, \tag{14}$$

where N is family size, in the loan size equation.

Model identification is due to classic parameter restrictions only. Exogenous variation in the cost of funds variable c helps identifying the crucial elasticity parameter ε. Variable $c = \ln(C(i, T))$ varies because i varies over time, but also because the loan term T varies at each period of time.[6] It is reasonable to assume that the cost of funds plays no direct role in the demand for loans, because borrowers care only about the loan price p. We add the constraint that the coefficient of ℓ in the price equation must be equal to $\lambda + \beta$. This is consistent with the French (and American) experience of lender underwriting behavior. On top of this, family size N is excluded from the price equation. Its inclusion in the loan amount equation helps identifying the coefficient of ℓ in the price equation.[7] The estimation methods used are 2SLS and 3SLS.

6. There are observations of up to 19 different loan terms from $T = 1$, to $T = 20$, at each period (observations being concentrated on $T = 10$, 15, and 20). There are also 6 different years.

7. Because of the constraint on ℓ's coefficient in the price equation, exclusion of N from the price equation is not necessary for identification. Our best justification for doing it is empirical: its inclusion does not yield a coefficient significantly different from zero.

3.2 **Model A: Estimation Results**

Tables 2-1 and 2-2 show the estimations of Model A. It has been estimated with the entire sample (all borrower categories pooled), and in each of the following four occupational status categories: Blue Collars (BC), White Collars (WC), Intermediate Professions (IP), and Executives (EX). Table 2-1 gives the results for the loan size, or quantity equation. Table 2-2 gives the corresponding results for the price equation. The quantity equation works quite well. There are categories of borrowers for which N is not significant, but this is of secondary importance. The downpayment variable a is not significant in sub-samples, except for executives, for which it has a negative sign (indicating substitutability of A and L in this category). The important facts are the significance of the price and wage variables. There is a remarkable ranking of the price-elasticity of demand estimates, which corresponds to the hierarchy of occupations: roughly, 2.02 for BCs, 2.03 for WCs, 2.39 for IPs, and 2.72 for EXs. Given the good precision of these estimates, 2.02 is significantly different from 2.39 and from 2.7. So it seems that the higher the occupational status, the higher the price-elasticity of demand. These differences, in turn, become the basis for differences in markups, and thus in the interest rates charged. A blue collar pays more for a loan than an executive, *ceteris paribus*, *i.e.*, even if they earn the same wage. The reason for this result might be that executives have better outside options than workers, when bargaining about loan conditions.

TABLE 2-1
Model A
Quantity equation

Dependent variable: logarithm of loan size (l) Estimated Coefficients (Student Ts)					
Explanatory variable	Entire sample(*)	Blue Collars	White Collars	Intermediate Professions	Executives
Price(p)	– 2.167 (– 40.52)	– 2.028 (– 21.81)	– 2.031 (– 19.91)	– 2.395 (– 14.67)	– 2.717 (– 9.381)
Constant	5.499 (28.47)	6.850 (20.44)	6.321 (17.38)	6.387 (9.935)	3.683 (3.685)
Household Size (N)	0.0253 (5.39)	0.03596 (6.009)	0.02945 (3.138)	0.01317 (0.7774)	– 0.02846 (– 1.012)
Wage (w)	0.1769 (10.11)	0.09564 (3.136)	0.1832 (5.714)	0.06340 (1.098)	0.4960 (5.523)
Downpayment (a)	0.07115 (8.94)	0.02170 (1.769)	0.005396 (0.3823)	0.03354 (1.4782)	– 0.1291 (– 3.338)
Adjusted R^2	0.445	0.212	0.234	0.226	0.253
Number of observations	2933	1180	909	363	159
Estimation Method	2SLS	3SLS	3SLS	3SLS	3SLS
Var(u)	– – –	0.1862	0.2545	0.4772	0.6772

TABLE 2-2
Model A
Price equation

Dependent variable: logarithm of price (p) Estimated Coefficients (Student Ts)					
Explanatory variable	Entire sample(*)	Blue Collars	White Collars	Intermediate Professions	Executives
Constant	– 0.3123 (– 21.63)	– 0.4652 (– 18.81)	– 0.4031 (– 15.51)	– 0.2168 (– 5.708)	– 0.3150 (– 5.267)
Cost of funds (c)	0.7945 (114.50)	0.7226 (47.37)	0.7630 (46.14)	0.8602 (34.69)	0.7655 (18.85)
Effort ratio ($l - w$)	– – –	– 0.001748 (– 0.3003)	0.007877 (1.247)	0.007686 (0.9692)	– 0.03119 (– 1.888)
Downpayment ratio ($a - l$)	– 0.01667 (– 12.00)	– 0.01534 (– 6.959)	– 0.01539 (– 6.734)	– 0.01317 (– 4.274)	– 0.01489 (– 3.434)
Adjusted R^2	0.846	0.772	0.807	0.875	0.833
Number of observations	2933	1180	909	363	159
Estimation Method	2SLS	3SLS	3SLS	3SLS	3SLS
Var(v)	– – –	0.004825	0.005817	0.005703	0.007691
Cov(u,v)	– – –	0.009271	0.007965	0.01951	0.04889
Correlation-coeff(u,v)	– – –	0.31	0.20	0.37	0.67

(*) The 2SLS estimation on the complete sample has been performed with the addition of 13 regional dummies indicating the BS (and therefore the region) where the loan was originated. We omit the coefficients and the t statistics associated with these dummy variables. Five of these dummies were significant in each equation. The complete sample contains other borrower categories, like farmers, retired persons, etc, which have not been studied separately. Regional dummies have not been included in the other estimations.

The price equation shows that the effort ratio is never significant (*i.e.*, $\widehat{\lambda} \simeq 0$), but, in contrast, the downpayment ratio is highly significant, with the expected negative coefficient (*i.e.*, $-\widehat{\beta} \simeq -0.016$): interest rates decrease when the downpayment ratio is high (or equivalently, when the initial Loan-to-Value is low). The coefficient of the cost variable, denoted δ, is significantly different from 1 (and from zero), around 0.8.

We should finally check that condition (11) holds. We accept that $\widehat{\lambda} = 0$. Using then $\widehat{\beta} = 0.016$, we find,

$$\widehat{\varepsilon} \simeq 2.16 > \frac{1}{1 - \widehat{\beta} - \widehat{\lambda}} \simeq 1.016,$$

a reassuring result. (This inequality holds in all sub-sample estimations too). To sum up, our simple model does not seem to be rejected by the data.

The order of magnitude of the relative price-cost margin implied by the estimated model can be computed with the help of expression (8) above. If we accept $\lambda = 0$, it is easy to check that the exact relative price-cost margin writes,

$$\frac{P - C}{P} = 1 - \pi \left(1 - \frac{1}{\varepsilon} - \beta\right). \tag{15}$$

To get the order of magnitude of π, we first evaluate μ for the entire sample. We get from estimations in Table 2,

$$\widehat{\mu} = e^{\widehat{\psi}} \left(1 - \frac{1}{\widehat{\varepsilon}} - \widehat{\beta}\right) \simeq (0.731)(1 - 0.465 - 0.015) \simeq 0.38,$$

and therefore $1/\widehat{\mu} \simeq 2.63$. From Table 1, we know that the mean value of A/L is approximately,

$$\widehat{A/L} = \left(\left(\widehat{\frac{A}{L+A}}\right)^{-1} - 1\right)^{-1} = \left(\frac{1}{0.172} - 1\right)^{-1} = 0.207.$$

To compute C, we take $T = 15$ and $i = 0.083$, so that, $C = C(0.083, 15)$ $\simeq 0.10$, and $C^{1-\delta} \simeq (0.1)^{0.2} \simeq 0.63$. This in turn yields,

$$\widehat{\pi} = (1/\widehat{\mu}) C^{1-\widehat{\delta}} (\widehat{A/L})^{\widehat{\beta}} \simeq (2.6)(0.63)(0.2)^{0.015} \simeq 1.606.$$

Using these values, we finally get the following reasonable figure,

$$\frac{P - C}{P} \simeq 0.165.$$

This figure, almost exactly one sixth, is very close to the empirical average of $(P - C)/P$ in the sample, which happens to be equal to 0.171. The reader must remember that this ratio is not expressed in interest rate terms, but in terms of constant repayment coefficients. The average value of $(r - i)/r$, a relative margin in terms of interest rates is equal to 0.285.

To show the hierarchy of markups, we compute the terms $\widehat{m} \equiv (\widehat{\beta} + 1/\widehat{\varepsilon})$ in each occupational status with the help of estimated coefficients. This yields, $\widehat{m} \simeq 0.50$ for blue and white collars, $\widehat{m} \simeq 0.43$, for intermediates, and $\widehat{m} \simeq 0.38$ for executives. These results already confirm that the higher the status, the lower the markup, and we can finally compute the implied values of the relative price-cost margin $(P - C)/P$. Computations give a margin of 20% for BC and WC, of 8.4% for IP, and of 0.4% for EX: a striking discrimination phenomenon. Empirical sub-sample averages of the relative margin are less extreme; we find 18.6% for BC and WC, $13{,}2\%$ for IP, and $11{,}8\%$ for EX.

We can also estimate the proper interest-rate elasticity of the demand for loans, using the formula,

$$-\frac{r}{L}\frac{dL}{dr} = -\left(\frac{P}{L}\frac{dL}{dP}\right)\left(\frac{r}{P}\frac{dP}{dr}\right).$$

In the neighborhood of $T = 15$ and $r = 0.1$, we find a value close to 1 for this elasticity. This means that if the interest rate, say, rises from 10% to 11%, the average loan size is expected to decrease by 10%.

The fact that we estimated the model in each occupational status subsample has some advantages. In particular, the price equation constants ψ are free to vary with status (as well as all other coefficients). The constants ψ might capture the effect of some status-specific default risk factors (*i.e.*, statistical discrimination based on status) and the effect of pure social prejudice (*i.e.*, what can be called "social discrimination"), if it exists. Due to the presence of these factors, the status-specific price-elasticity estimates could have been biased if these constant terms had been constrained to be equal accross occupation groups. Finally the model parameters are remarkably stable, especially if one compares BCs and WCs: this is reassuring.

3.3 Model B

We have also studied a variant of Model A, using the idea of demand for housing instead of demand for loans. In Model B, we specify the demand for housing as follows,

$$H = L + A \equiv \Phi(N)W^{\gamma}A^{\sigma}P^{-\varepsilon}. \tag{16}$$

We then accordingly modify the pricing equation as follows,

$$P = \Psi C^{\delta}\left(\frac{A}{H}\right)^{\beta}. \tag{17}$$

The effort ratio has been suppressed from the equation (λ was not significantly different from 0). This formulation is slightly ad hoc, because it cannot be justified as a monopoly pricing equation as easily as Model A's pricing equation. We have nevertheless estimated this model. Taking logs and adding random disturbances (u, v) again, the econometric specification of Model B writes,

$$h = \varphi_0 + \varphi_1 N + \gamma w + \sigma a - \varepsilon p + u, \tag{18}$$

$$p = \psi + \delta c - \beta(a - h) + v. \tag{19}$$

This model is overidentified and can be estimated by 2SLS and 3SLS exactly as Model A. Variable $h = \ln(L + A)$ now replaces $\ell = \ln(L)$.

Tables 3-1 and 3-2 show the results. Estimation results are better than that of Model A above. No wonder, the downpayment variable is now significant in the quantity equation, with a positive coefficient. Family size works better, except for executives. The adjusted R^2 is remarkably higher than the corresponding value for Model A. These results permit one to appreciate the robustness of the results obtained with Model A. This variant shows the same hierarchy of price-elasticity estimates. The values of ε are now much lower, around 1.6, instead of 2.15, but they do not measure the same thing. To see this, note that if the demand for loans is isoelastic and writes $L = KP^{-\varepsilon}$, then, by definition, $h = \ln(A + KP^{-\varepsilon})$. Thus in model B, instead of ε, we in fact estimate the parameter,

$$\varepsilon' = -\frac{\partial h}{\partial p} = -\frac{\partial \ln(L + A)}{\partial \ln(P)} = \frac{L}{L + A}\varepsilon < \varepsilon.$$

It is therefore not surprising to obtain smaller elasticities (in absolute value) with Model B. Note in addition that the average value of $A/(L + A)$ being equal to 0.172 (see Table 1), we get $L/(L + A) \simeq 0.828$; and using this value, we find,

$$\varepsilon' = \varepsilon L/(L + A) \simeq (2.15).(0.828) \simeq 1.78.$$

This result is consistent with the theory expressed by Model A: Model B should provide 20% smaller estimates of the price-elasticity. Again, IPs and EXs seem to have higher price-elasticities of their demand for loans than BCs and WCs, but all differences are not highly significant. White collars (or blue collars) are significantly different from executives (or IPs), however.

TABLE 3-1
Model B
Quantity equation

Dependent variable: logarithm of house size (h)
Estimated Coefficients (Student Ts)

Explanatory variable	Entire sample(*)	Blue Collars	White Collars	Intermediate Professions	Executives
Price(p)	– 1.6245 (– 37.93)	– 1.551 (– 20.75)	– 1.347 (– 16.27)	– 1.629 (– 12.74)	– 1.648 (– 7.240)
Constant	4.990 (32.27)	6.239 (23.15)	6.222 (21.09)	5.918 (11.85)	4.752 (6.117)
Household Size (N)	0.02072 (5.521)	0.0280 (5.823)	0.02397 (3.145)	0.01487 (1.091)	– 0.01631 (– 0.7071)
Wage (w)	0.1585 (11.32)	0.102 (4.187)	0.1486 (5.713)	0.05215 (1.169)	0.3275 (4.726)
Downpayment (a)	0.2642 (41.48)	0.193 (19.60)	0.2043 (17.84)	0.2652 (14.92)	0.1533 (5.040)
Adjusted R^2	0.571	0.421	0.401	0.477	0.350
Number of observations	2933	1180	909	363	159
Estimation Method	2SLS	3SLS	3SLS	3SLS	3SLS
Var(u)	– – –	0.1155	0.1595	0.2684	0.3685

TABLE 3-2
Model B
Price equation

Dependent variable: logarithm of price (p) Estimated Coefficients (Student Ts)					
Explanatory variable	Entire sample(*)	Blue Collars	White Collars	Intermediate Professions	Executives
Constant	– 0.3287 (– 23.05)	– 0.4787 (– 19.39)	– 0.4280 (– 16.73)	– 0.2454 (– 6.850)	– 0.3191 (– 5.409)
Cost of funds (c)	0.7928 (115.01)	0.7239 (62.82)	0.7452 (61.26)	0.8408 (47.70)	0.8159 (29.40)
Downpayment ratio ($a - h$)	– 0.02091 (– 11.94)	– 0.01862 (– 6.96)	– 0.01931 (– 6.79)	– 0.01671 (– 4.170)	– 0.01586 (– 2.916)
Adjusted R^2	0.846	0.770	0.807	0.869	0.852
Number of observations	2933	1180	909	363	159
Estimation Method	2SLS	3SLS	3SLS	3SLS	3SLS
Var(v)	– – –	0.004868	0.005813	0.005972	0.006794
Cov(u,v)	– – –	0.005797	0.005403	0.01218	0.01619
Correlation-coeff(u,v)	– – –	0.24	0.18	0.30	0.32

(*) The 2SLS estimation on the complete sample has been performed with the addition of 13 regional dummies indicating the BS (and therefore the region) where the loan was originated. We omit the coefficients and the t statistics associated with these dummy variables. Five of these dummies were significant in each equation. The complete sample contains other borrower categories, like farmers, retired persons, etc, which have not been studied separately. Regional dummies have not been included in the other estimations.

4 A Nonlinear Market Equilibrium Model

The differences in price-elasticity of demand observed above seem to be a basis for some discriminatory pricing practice of the banker. To perform a different test of the presence of such interest-rate discrimination, we now construct a simple structural model of competition in mortgage loan contracts, in which the borrower's preferences are explicitly specified.

4.1 Basic assumptions

Assume that each borrower is characterized by extreme impatience, and is severely liquidity constrained. Let ρ be the price of housing per square meter, which is assumed constant. The consumption of housing is simply defined as

$$H = \frac{L + A}{\rho}.$$

We adopt a continuous-time specification of the model. Let r now be the continuous-time interest rate. It can be computed easily from the yearly interest rate R with the formula, $R = e^r - 1$, when time T varies continuously with the year as a unit. With these assumptions, the instantaneous, constant amortizing repayment of a one euro loan over T years becomes,[8]

(20) $$P = P(r,T) = \frac{r}{1 - e^{-rT}}.$$

Assume that the instantaneous utility of a borrower of preference-type (α,γ) depends on consumption X, and housing H, and writes,

(21) $$U(X,H) = X + \gamma\frac{H^{\alpha}}{\alpha}.$$

The price-elasticity of the demand for housing is defined as,

(22) $$\varepsilon = \frac{1}{1 - \alpha}.$$

The borrower's budget constraint at each period is simply $X + PL = W$, where L is the amount lent and W is instantaneous income (wage). A mortgage loan is fully defined by (r,T,L), that is, by (P,L). Given characteristics (A,W), there is a one-to-one transformation taking the "contract space" to the "consumption-housing space", that is, from (P,L) to (X,H). More precisely, the transformation is $X = W - PL$ and $H = (1/\rho)(L + A)$.

We now compute the expected profit of a loan to a given observable type of borrower. We first add the "risk-class" θ of the borrower to preference parameters (α,γ). Assume that each worker-borrower loses his(her) job (and defaults) at a random time t, which is exponentially distributed with parameter $1/\theta$. In other words, we assume that the probability of defaulting between time 0 and time t is $\Pr(\tau \leqslant t) = 1 - e^{-\theta t}$. With this assumption, we can compute the expected benefit of a banker from a loan to a borrower in risk-class θ. We assume for simplicity that mortgage foreclosure yields zero revenues when default occurs at any time before the term T. Conditional on default occurring at time $t \leqslant T$, the present value of a loan is therefore,

(23) $$b(t) = -L + LP(r,T)\int_0^t e^{-i\tau}d\tau,$$

8. Remark that, by definition of $P(r,T)$, we must have, $1 = \int_0^T P(r,T)e^{-rt}dt$.

where i is the banker's cost of funds and discount rate. Expected benefit B then writes,

$$B = \int_0^T \theta e^{-\theta t} b(t)dt + e^{-\theta T} b(T),$$

and, after straightforward integration, we get the simple formula,

$$B = -L + L\frac{P(r,T)}{P(i+\theta,T)}. \tag{24}$$

If we now assume that the management costs of a loan file are negligible for the bank, the zero profit condition, *i.e.*, $B = 0$, simply writes,

$$P = P(i+\theta,T), \tag{25}$$

where $P = P(r,T)$ is the "price" charged for the loan. To avoid possible confusion, we again rename the function as follows, $C(i+\theta,T) \equiv P(i+\theta,T)$, and use C when this term plays the role of a "cost" for the banker.

Using the one-to-one change of variable defined above, we can rewrite expected profit as a function of (X,H). We get,

$$B = -\rho H + A + \frac{W - X}{C(i+\theta,T)}, \tag{26}$$

which is a linear function of X and H.

4.2 Competition in contracts

To specify the model fully, a market equilibrium notion must be defined. Assume for simplicity that competition in contracts prevails among banks, with a condition of free entry on the banking market. Free entry drives profits to zero for each observable borrower type, *i.e.*, for each risk class and each term T, we have $P = C(i+\theta,T)$. In addition, the contract offered to each type of borrower, and each T, maximizes U under the zero-profit constraint (*i.e.*, no banker can enter the market and offer a more advantageous contract (P,L) to the same type of borrowers). It follows that in equilibrium, contracts are such that the borrower's indifference curve is tangent to the banker's zero profit line in the consumption-housing plane (X,H). Formally, we must have in equilibrium,

$$\frac{\partial B/\partial X}{\partial B/\partial H} = \frac{\partial U/\partial X}{\partial U/\partial H}. \tag{27}$$

With our specifications, this equilibrium condition can be rewritten as follows,

$$\frac{1}{\rho C(\theta + i,T)} = \frac{H^{1/\varepsilon}}{\gamma},$$

or, finally,

$$H = \left(\frac{\gamma}{\rho C(\theta + i,T)}\right)^{\varepsilon}. \tag{28}$$

This relation determines equilibrium house size as a function of preference parameters (ε,γ), and of the risk-premium θ. To estimate the model, we will let these parameters vary with borrower characteristics.

4.3 Econometric Estimation

Our theory suggests a bivariate econometric model, determining H and P simultaneously. To obtain the model, we simply assume that there are multiplicative, log-normal perturbations affecting the zero-profit and the tangency conditions (25) and (28) derived above. Taking logarithms in (25) and (28) and adding perturbations yields the nonlinear system,

$$h = -\varepsilon \ln(\rho) + \varepsilon \ln(\gamma) - \varepsilon \ln(C(\theta + i,T)) + u, \tag{29}$$

$$p = \ln(C(\theta + i,T)) + v, \tag{30}$$

where $h = \ln(H)$, and $p = \ln(P(r,T))$.

To specify the model completely, we assume, first, that (u,v) is a normally distributed vector with mean vector 0 and variance-covariance matrix Ω; second, that γ and θ are functions of observable borrower characteristics. More precisely, we set,

$$\ln(\gamma) = \gamma_0 + \gamma_1 w + \gamma_2 y + \gamma_3 n, \tag{31}$$

where, $w = \ln(W)$, $n = \ln(N)$, $y = \ln(Y)$ and Y is the age of the borrower. To specify the risk-premium θ, we set,

$$\begin{aligned} \theta = \exp\Bigg[& \theta_0 + \theta_1 w + \theta_2 w \cdot (EX) \\ & + \theta_3 w \cdot (IP) + \theta_4 w \cdot (WC) + \theta_5 \ln\left(\frac{L+A}{A}\right)\Bigg]. \end{aligned} \tag{32}$$

In the above expression of θ, EX, IP and WC are occupational status dummies for executives, intermediate professions and white collars, respectively, interacting with the wage w. Blue collars are the reference group.

Third, we let the price-elasticity of demand for housing ε vary with occupational status, adding dummies EX, IP and WC. Blue collars are again the reference group.

Fourth, given that we lack data about the real estate prices ρ we simply "calibrate" ρ by choosing a reasonable (constant) value (*i.e.*, $\rho = 42.8$ euros per square meter and per year). Note that this choice is harmless, its only effect being to change the magnitude of the constant γ_0 in the expression of $\ln(\gamma)$.

The model is given by (29), (30), (31) and (32). Remark first that the model has a special recursive structure: p does not intervene in (29), but only its expected value $\ln(C(\theta + i, T))$. Now, this latter function depends on $\ln(L + A)$ and thus on h through (32), the risk-premium function. It follows that (29) is in fact a nonlinear relation involving a single endogenous variable, *i.e.*, h only. It is important to note that this loan size equation is not a demand function: strictly speaking, it describes the equilibrium "menu of contracts". The model permits the identification of ε anyway.

To understand how this model is identified, a linear approximation of $\ln(C(\theta + i, T))$ can be used. If it is assumed, for instance, that $\ln(C(\theta + i, T))$ is a linear function of i and $\ln(\theta)$, the model becomes a standard linear simultaneous equation system and it becomes straightforward to check that it is overidentified, identification being entirely due to classic parameter restrictions. Next, the function $\ln(C(\theta + i, T))$ is a strictly increasing function of θ, so that the nonlinear model is well behaved (the value of h is determined uniquely by (29) and thus, the value of p is also uniquely determined by (30)). There should be (and in fact there were) no identification problems.

The model has been estimated by the full information maximum likelihood method, $\ln(P)$ and $\ln(L + A)$ being the endogenous variables. Since $\ln(L + A)$ appears in the expression of θ, the log-likelihood function is less simple than it could seem at first glance: there are non-trivial Jacobian terms.

Estimation results are reported in Table 4. The price-elasticity ε is again of the order of magnitude of 2.17. There are some significant differences of this elasticity from one group to the other; essentially, IPs and executives are significantly different from BCs (BCs are the reference group, with $\varepsilon = 2.116$). True, blue collars are not significantly different from white collars, but the hierarchy of elasticity values appears again. A number of other things are captured: the effect of wage and of the Loan-to-Value ratio on the risk-class parameter θ; and the effect of wage and age on parameter γ. It permits an analysis of the impact of observable borrower characteristics on both the interest rate and the size of the loan simultaneously, in a consistent way. The values of θ obtained are too high to reflect simply risk premia, they are more realistically interpretable as monopolistic markups. We have attempted to deal with this difficulty in a companion paper (GARY-BOBO and LARRIBEAU [2004]).

5 Links with the literature on racial discrimination in mortgage lending

Discrimination in credit markets has recently attracted considerable attention in the US, and the question of deciding whether or not – and why –

TABLE 4
Nonlinear Model

Parameter	Estimation	T-Ratio	P-Value
Epsilon (ε)			
– Constant	2.1166	29.908	0.0000
– Executive	0.0600	3.146	0.0008
– Intermediate	0.0309	2.223	0.0131
– White Collar	0.0065	0.645	0.2594
Gamma			
– Constant (γ_0)	5.2850	39.677	0.0000
– Wage (γ_1)	0.2061	15.588	0.0000
– Age (γ_2)	0.1017	5.662	0.0000
– Household Size (γ_3)	0.0067	0.848	0.1981
Theta			
– Constant (θ_0)	– 2.8077	– 9.330	0.0000
– Wage (θ_1)	– 0.0749	– 2.858	0.0021
– Wage*Executive (θ_2)	– 0.0049	– 1.065	0.1434
– Wage*Intermediate (θ_3)	0.0011	0.386	0.3498
– Wage*WhiteCollar (θ_4)	0.0072	3.980	0.0000
– Inverse Downpayment Ratio (θ_5)	0.1945	19.081	0.0000

Mean of Theta for		Epsilon :	
– Executives	0.0330	– Executives	2.1766
– Intermediates	0.0367	– Intermediates	2.1475
– White Collars	0.0425	– White Collars	2.1231
– Blue Collars	0.0398	– Blue Collars	2.1166

Estimation method: Maximum Likelihood

Estimated Covariance matrix of errors :

$$\Omega = \begin{pmatrix} 0.2083 & -0.0087 \\ -0.0087 & 0.0110 \end{pmatrix}$$

Correlation Coefficient = – 0.1814

Number of observations = 2610

Mean Log-Likelihood = 0.305248

Empirical variances :
$\mathrm{Var}(\ln(L + A)) = 0.2958$
$\mathrm{Var}(\ln(P)) = 0.0286$

lenders discriminate against minority groups is a hotly debated topic among economists. The importance of the question is amplified by the fact that racial discrimination in mortgage lending has been made illegal by the Equal Credit Opportunity Act of 1974, and by the availability of new data sources, allowing for new econometric tests. The recent literature on this question is mostly empirical, and concentrated on racial or sexual discrimination problems. Empirical studies of discrimination in mortgage lending have developed with the debate triggered by the contributions of SHAFER and LADD [1981] and

MUNNELL *et al.* [1996]. Several other important contributions to this literature are commented in LADD'S [1998] survey article, and in the recent book of ROSS and YINGER [2002]. Much of the published work on discrimination in the mortgage market, as far as we know, has been devoted to the study of default and of credit denial rates. In contrast, the model presented here aims at explaining the structure of *accepted* loan applications and determines loan sizes and interest rates simultaneously.

A difficult problem in most empirical studies is to detect the presence of discrimination in the sense of G. BECKER[9] [1957], and to distinguish it from the effects of default risk and of statistical discrimination[10]. Another difficulty is the need to separate the effects of prejudice from that of market power in the formation of interest rates. Competition should tend to eliminate discrimination in the sense of Becker, since prejudiced lenders would lose business in favor of the unprejudiced. It follows that market power and discrimination in this sense must also be closely interrelated.

If the borrowers' preference parameters are correlated with their social, racial or ethnic group, because individual "preferences" depend on a group's particular economic conditions and environment, and if market power is present, then, standard third-degree price discrimination can become the explanation for differential treatment, without necessarily reflecting the presence of prejudice. Third-degree price discrimination, based on differences in price-elasticities of demand can then be viewed as a special type of statistical discrimination, in which observable group identity is used as a signal for different preferences. This type of approach, however, is not without its dangers, which would be to attribute the bulk of observed differences in treatment to taste differences correlated with race. On the role of concentration in the explanation of interest rates, see CAVALUZZO and CAVALUZZO [1998]. On the empirical relationship between concentration and discrimination, see BERKOVEC *et al.* [1998], and ROSS [1996]. See also ROSS [2000], on tests for prejudice-based discrimination.

Our models could be estimated with other data sets, including observations of race. It is sufficient to have a sample of accepted loan files, with at least an observation of the loan rate, loan amount, downpayment, term and conditions, starting date, and some observations of borrower and house characteristics. More could be done with richer observations of the borrower and his (her) family, with informations on the house purchased, on house prices and characteristics of the neighborhood. Informations on the borrower's job and employer (seniority, size of the firm, industrial sector, etc.) along with the borrower's qualifications and years of schooling would be of great value to estimate his (her) probability of loosing his (her) job. Our models can be easily adapted to exploit these observations. In model A, the pricing and demand functions (12) and (13) could depend on a much longer list of exoge-

9. A lender, or seller, is said to discriminate in the sense of Becker, if she is ready to forego profits just because of her prejudices. This form of discrimination is based on a particular "taste for discrimination" of the sellers, and is not usually considered by standard I.O. theories of price discrimination.

10. A lender might treat a minority group differently, because racial or ethnic characteristics are correlated with some variables, important in the determination of creditworthiness and default risk, and which remain unobserved. This latter form of behavior is called *statistical discrimination* in the sense of ARROW [1973] and PHELPS [1972].

nous variables. In the demand equation, the price-elasticity can be estimated more finely in borrower subgroups, if price is interacting with subgroup dummies. The same can be said of the nonlinear model, which is probably more adapted to competitive credit market environments (*i.e.*, the US). Both models could then be used to test for the presence of racial discrimination. If race has a significant coefficient in the price (risk-premium) function, then, the richer the set of controls for default risk are also included in the regressions, the stronger the suspicion that discrimination à la Becker is present.

6 Conclusion

To conclude, in the present contribution, we proposed several variants of a model of the mortgage lending market. The models can be used to test for the presence of market power and discrimination, using information on accepted loan applications only. They rest on the idea of discrimination by the lender, based on observable attributes of the borrower. They explain the interest rate and the loan size of accepted loan applications simultaneously. We study a market equilibrium variant and a monopolistic variant of the model. Both variants provide good estimates of the interest-rate elasticity of demand for mortgage loans.

The model has been estimated with a sample of loan files originating from branches of a French mortgage lender. Part of the observed differences in interest rates must be attributed to differences in structural preference parameters, since blue-collar workers have a significantly smaller price-elasticity of demand for housing than executives. These differences reflect the different situations and environments of socio-economic groups, and become a basis for differential treatment by bankers. ▼

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