bedrooms, paved road, and median family income. While the individual effects cannot be identified, the insignificance of important structural attributes is consistent with equalization of marginal attribute prices.

Finally, there is some evidence of heterogeneity in the data. As would be expected with heterogeneity, use of the Weibull-Gamma hazard increases the point estimates of the duration elasticities. In addition, a likelihood ratio test finds the estimated variance of the Gamma distribution is significantly greater than 0 at the 20% level. It is interesting that heterogeneity does not appear to be as strong in these data as that typically found in labor market applications of the hazard methodology. Perhaps these results reflect the usefulness of asking prices and occupancy status as summary measures of variation in seller characteristics that influence the choice of reservation prices and thus probability of sale.

IV. Conclusions

Empirical analysis of housing markets has focused on the heterogeneous nature of housing and the resulting consequences for price determination. This study employs a search theoretic approach to examine the relationship between probability of sale and market duration in the housing market. Stronger incentives for a diminishing reservation price are identified for vacant houses, suggesting a higher rate of time dependence for these houses relative to occupied houses. This hypothesis is supported by the empirical evidence presented in this paper, where vacant houses are found to exhibit positive duration dependence, while little evidence of duration dependence is found for occupied houses. Finally, while limited by the fact that only net effects can be identified, the insignificance of important structural attributes in explaining cross-section variation in probability of sale suggests that the introduction of price uncertainty does not diminish the importance of hedonic prices in the determination of attribute choice, or the resulting tendency of these prices to equalize across buyers and sellers.

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LOAN DEMAND: AN EMPIRICAL ANALYSIS USING MICRO DATA

John C. Ham and Arie Melnik*

Abstract—This paper analyzes data on loan commitment contracts collected from a sample of non-financial firms. We first discuss general features of the contracts and present sample summary statistics. We then estimate an equation describing firm borrowing behavior. We use both regression analysis and an estimation approach that, under certain assumptions, allows for censored data, since 20% of the firms used their entire credit allocation at least once.

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1. Introduction

The demand for business loans has been investigated thoroughly in the past (e.g., Goldfeld (1969) and Wood (1975)). More recently, Hicks (1980) examined factors that influence aggregate business loans and estimated a detailed loan demand function. However, these studies (and others) used aggregate data. Our purpose in this paper is to utilize micro data to estimate the determinants of business borrowing behavior. Since most businesses are granted loans through loan commitment contracts we focus on borrowing against existing credit lines.

We collected loan commitment data from a sample of non-financial firms. We use our data in two ways to provide additional knowledge of the corporate credit market. First, we discuss a number of general features
of these contracts and present summary statistics for
our sample. Second, we use our data to estimate the
determinants of a firm's borrowing behavior. In the
estimation we face a problem of censored data since
20% of the firms in our sample used their entire credit
allocation at least once. We adopt an estimation scheme
which addresses this problem.

The outline of the paper is as follows. In section II we
provide a discussion of the institutional features of the
credit contracts from our sample. In section III we
describe a statistical model for estimating a credit
demand equation based on micro data including
censored observations. The results of implementing this
procedure are presented in section IV.

II. Institutional Features of Loan Contracts

In a loan commitment contract, a bank (or a group of
banks) agrees to meet future customer loan needs, up to
some pre-specified amount.1 This predetermined future
amount is called the "commitment." The contract may
be viewed as a contingent asset of the borrower and a
contingent liability of the bank. The future lending is
made at a fixed pricing formula. In most cases, the
interest rate is linked to the prime rate or to some other
agreed upon market-determined interest rate; it is
generally not fixed over the life of the contract. Specifi-
cally, the pricing of the contract is composed of two
elements: a general market (or market linked) rate and
a customer-specific risk premium. In addition, the
contract itself entails a cost called the "commitment
fee" which the customer pays for the privilege of having
standby credit. This fee is usually accrued continuously
on the unused balance of the credit line. The typical
credit line contract contains, in addition to the commit-
ment size and the price to be paid for possible future
loans, restrictive covenants (e.g., collateral require-
ments, limited dividends, etc.). These are designed either
to monitor the behavior of the borrower or to enhance
the safety of the contract.

Our sample of corporate borrowers provides some
information on the features of loan contracts. The data
were collected from responses of corporate treasurers to
a questionnaire which was mailed to a sample of U.S.
non-financial corporations. The mailing list containing
420 borrowers was selected at random from the Stan-
ard and Poors' register for corporations and consti-
tuted 3% of a population of about 14,000 companies.
Only 132 firms (31% of the sample) provided answers. A

number of firms gave only partial information. Answers
about maximum take-downs were provided by 90 firms
and they are used in the estimation below.2

In our sample about 40% of the contracts were signed
between one customer and one bank, 24% involved one
borrower and two banks and in 15% of the cases three
banks jointly appeared on the lending side. In the
remaining 21% of the contracts four or more banks were
joint parties to the contract. The number of banks on
the lending side appears to be associated with the size
of the loan. This association may be explained either by
legal lending constraints or by the desire to spread the
lending risk.

The firms paid, on average, a risk premium of 0.837%
above prime, with a range from 0% to 3.5%. Fischer
(1982) mentioned the practice of below prime lending.
None of the contracts that we surveyed contained a
below prime rate. However, it should be noted that,
according to Goldberg (1981), below prime loans were
granted occasionally by the largest banks to their largest
customers. Our sample does not contain many unusually
large borrowers and hence does not capture such cases.
The average commitment fee in our sample was 0.4% of
the undrawn amount. It ranged from 0.1% to 0.6%.

Only 55% of these contracts contained a compensat-
ing balances requirement. When such requirements were
imposed they ranged from 5% to 20% of total borrow-
ing, with an average of 11%. Only 20% of the contracts
contained a formal collateral requirement.

An interesting feature of the contract is its time span.
The mean contract length was approximately 28
months. The minimum contract length was 6 months
and the maximum length was 84 months. The length of
the contract may reflect a formalization of the bank-
customer relationship (Hodgman (1963), Hester (1979))
and may be explained by the desire—on the part of
both parties—to economize on loan transaction costs.

A very important feature of the credit line contract is
the quantity setting clause. This commitment obligation
places an upper limit on the customer's borrowing. In
our sample the mean commitment size was $7.64 mil-
lion. Within this limit, the borrower may set future loan
size unilaterally at the agreed upon rate. In our sample
the average maximum actual take-down was $5.1 mil-
lion. However, approximately 20% of the firms in our
sample reached the maximum limit of their commit-
ment size. This raises the possibility that these firms
face credit rationing, in the sense that they cannot
borrow as much as they want at the going rate of
interest.3

---

1According to the Federal Reserve surveys of loans, the
majority of industrial loans granted by commercial banks in
the United States are made under loan commitment contracts.
This phenomenon is even more pronounced in the case of large
(greater than $0.5 million) loans, over 70% of which are made
under some form of credit line arrangement.

2In view of the partial response rate, our sample may suffer
from a response bias in that firms that answered our inquiry
could differ from those that did not.

3Models of credit rationing date back to the sixties (Hodg-
man (1963), Freimer and Gordon (1965), Jaffee and Modigliani
In determining whether the firms that exhaust their credit line in our sample can properly be described as rationed, it is important to ascertain the options open to the firm once it reaches its credit limit. If such a firm is unable to renegotiate its contract and is unable to borrow additional funds from alternative sources, it is sensible to describe the firm as rationed. However, if either of these options is available to the firm, with additional borrowing requiring a higher interest rate, it is more appropriate to describe the situation as one where the firm faces a nonlinear price schedule.

Our data set does not allow us to determine which description is more appropriate. However, our sample indicates that rationing is a potential problem since the fraction of firms facing such rationing could be as high as 20%. Moreover, in either of the above cases, firms that reached their credit limit wanted to increase their borrowing at the interest rate specified in the contract. This phenomenon should be taken into account in the empirical estimation of the loan demand equation. We now consider this issue.

III. Estimating a Loan Demand Equation

Our equation for the current desired borrowing by each firm, \( B_i^d \), takes the form

\[
B_i^d = B_i(r^*, \delta, T) + \epsilon_i. \tag{1}
\]

In (1) \( r^* \) is the net interest rate faced by borrower \( i \), \( \delta_i \) is a scale variable for the size of the borrowing firm, and \( T \) represents aggregate economic shocks which may affect the firm's desired borrowing. The variable \( T \) is not firm-specific but is determined by aggregate credit conditions. The error term \( \epsilon_i \) is assumed to be homoskedastic and independent of \( r^*, \delta_i \), and \( T \). For simplicity, we assume that (1) may be written in linear form as

\[
B_i^d = X_i'\gamma + \epsilon_i, \tag{2}
\]

where \( X_i \) is a vector of independent variables from (1) and \( \gamma \) is a vector of parameters. If the value of \( B_i^d \) was known for each firm then we could use regression techniques to estimate (2). However, for firms that reach their credit limit, we only observe the credit commitment \( C_i \). For such firms, the level of desired borrowing is censored, and we only know that \( B_i^d > C_i \). One possible approach in estimating \( \gamma \) is to ignore the censoring, replace \( B_i^d \) by \( C_i \) in (2) for firms that reach their limit, and then use least squares to estimate (2). However, this will lead to biased estimates of \( \gamma \) if any of the \( X_i \) variables are correlated with the difference between desired and actual borrowing. To see this, rewrite (2) as

\[
B_i = X_i'\gamma + \epsilon_i + (B_i - B_i^d) \tag{3}
\]

where \( B_i \) represents actual borrowing and \( B_i - B_i^d = C_i - B_i^d < 0 \) for the censored firms and zero otherwise. In this case, the expected value of the least squares estimate of \( \gamma_i \) will equal

\[
E(\hat{\gamma}_i) = \gamma_i + \delta E(B_i - B_i^d)/\delta X_i. \tag{4}
\]

For example, suppose that firm size is a determinant of borrowing behavior, and that larger firms are less likely to exceed their credit limit. Then the second term in (4) will be positive, and least squares will overestimate the effect of firm size on desired borrowing.

As an alternative to least squares estimation, we consider parameter estimates based on a simple statistical model which allows for censoring

\[
B_i = X_i'\gamma + \epsilon_i \quad \text{if} \quad B_i^d < C_i, \tag{5}
\]

\[
B_i = C_i \quad \text{if} \quad B_i^d \geq C_i.
\]

We assume that \( \epsilon_i \sim i.i.d. N(0, \sigma^2) \) and that \( C_i \) is distributed independently of \( \epsilon_i \). This leads to a variable limit Tobit model. Letting \( R \) denote the set of all borrowers who were censored and \( N \) denote the set of all remaining borrowers, the likelihood function is

\[
L = \prod_{i \in R} (1 - F(Z_i)) \prod_{i \in N} \frac{f(Z_i)}{\sigma} \tag{6}
\]

6 The disequilibrium literature is based on the observation that, in the presence of credit rationing and in the absence of information on which borrowers are rationed, all observations are censored in the sense that only the minimum of borrow demand and supply (i.e., the credit line) is observed. Sealy (1978) estimates such a model for the aggregate loan market, while Avery (1981) estimates a disequilibrium model for the borrowing behavior of individual consumers using micro data. Our work differs from Sealy in that we use micro data while he uses aggregate data, and differs from Avery in that we analyze firm behavior while he investigates consumer borrowing. Our work differs from both of these studies since we have sample separation information and can determine which observations are censored. Note, however, that we make an additional independence assumption.
where
\[ Z_i = \left( \frac{B_i - X'_i \gamma}{\sigma} \right) \]
(not including \( B_i = C_i \) if \( i \in R \)). To estimate \( \gamma \) and \( \sigma \) we maximize the log-likelihood function
\[ L^* = \sum_{i \in R} \ln(1 - F(Z_i)) + \sum_{i \in N} [-\ln \sigma + \ln f(Z_i)] \] (7)

IV. Empirical Results

We estimate the parameters of the desired borrowing equation (2) both by least squares and by maximizing the log-likelihood function (7). Since the data consist of the cross-section of firms from our survey, we assume that (2) is valid for the firm's maximum borrowing behavior. Our choice of independent variables is based on previous credit demand studies, and consists of the following variables: sales, interest cost, the level of borrowed reserves and a dummy variable coded 1 if the borrower placed collateral with the bank. (See table 1 for the mean values of the variables used in estimation.)

A few words of explanation are in order regarding the variables used. We chose the value of total sales in the preceding period as a scale variable. The interest rate variable is also firm-specific. The loan rate, as stated in the contract, contains the prime rate plus a customer-specific risk premium. For some firms the base rate is just the prime rate (this is the rate charged to the best customers); for others an add-on premium is attached.\(^7\)

Following contract specifications for each firm we compute the cost of borrowing by taking the prime rate in the month when the reported credit take-down took place and adding to it the risk premium, where appropriate.\(^8\) We then subtract the loan commitment fee from the resulting total since this component is most commonly paid only on the unused portion of the credit line.

We use borrowed reserves to proxy fluctuations in the economy which affect the firm’s demand for credit. In fact, a major reason for signing a loan commitment contract and paying a non-trivial commitment fee is to ensure an availability of credit in states of nature when tightness occurs. An increase in the level of this variable reflects a general increase in monetary tightness. Finally, a dummy variable for collateral is added to reflect inventory financing needs. The most common collaterals are inventories of finished goods and receivables obtained from customers in the course of business. Therefore, collateral may serve as a proxy for inventory financing needs which in turn are often used in explaining business loan demand. (See, for example, Hicks (1980) and the references cited there.)

In sum we expect a positive relationship between the maximum borrowing and total sales. The relation between interest cost and borrowing should be negative. The coefficients of borrowed reserves and of collateral are expected to be positive.

The least squares and maximum likelihood estimates are presented in table 1. The parameter estimates from the two estimation methods are qualitatively similar and have the expected signs. Using the maximum likelihood estimates, the elasticity of desired borrowing with respect to an increase in the premium at the sample means is estimated at 0.325. The elasticity of desired borrowing with respect to an increase in the prime is -5.23. Changes in the prime rate are likely to be accompanied by changes in borrowed reserves. Therefore, we also estimated the elasticity of demand with respect to the prime while not holding borrowed reserves constant. The resulting elasticity is estimated at -1.46. The elasticity of desired borrowing with respect to total sales is approximately 0.715.

We should note that our measure of the net interest cost restricts the coefficients on the prime, the premium and (minus) the commitment fee to be the same. Alternatively, conditional on the net borrowing cost, our specification does not allow the premium and commitment fee to play a separate role in the borrowing equation. This may be overly restrictive, particularly in the case of the risk premium. Firms in different risk classes (and thus with different risk premiums) may differ in their borrowing behavior.\(^9\) To investigate this issue, we first added both the risk premium and the commitment fee to our specification of the borrowing equation given in table 1. When we tested the hypothesis that the coefficients on these additional variables were jointly zero, we obtained a likelihood ratio statistic

\(^7\) The prime plus convention indicates that the customer specific risk premium is viewed by the bank as independent of the general movement of interest rates. The prime itself is viewed by both borrowers and lenders as reflecting an exogenous (non-customer specific) cost component of bank funds. The definition of the prime rate itself is not identical in all contracts but the variations are rather small. In 7 contracts the risk premium was a multiple (such as 0.1) of the prime. In the empirical work we calculated the premium for these firms in two ways. First, we calculated the average premium over the period covered by their contract. Second, we calculated the risk premium in the month of maximum borrowing. The results reported in section IV are based on the average premium, but using the alternative measure for these firms produced essentially identical results.

\(^8\) One may argue that while it may be reasonable to treat the prime as independent of \( \epsilon_t \) in (2), it is not reasonable to treat the premium or commitment fee as exogenous when estimating this equation. To investigate this issue, we replaced the firm’s net interest rate by a fitted value based on the prime and the other variables in (2). The least squares and maximum likelihood estimates based on this fitted value were virtually identical to their counterparts based on the actual value of the net interest cost.

\(^9\) This point was noted by a referee.
THE REVIEW OF ECONOMICS AND STATISTICS

Table 1.—Estimates of the Demand for Credit Function

<table>
<thead>
<tr>
<th>Coefficient of</th>
<th>Mean Value</th>
<th>Least Squares</th>
<th>Maximum Likelihood</th>
</tr>
</thead>
<tbody>
<tr>
<td>Constant</td>
<td>—</td>
<td>12.342</td>
<td>18.059</td>
</tr>
<tr>
<td></td>
<td></td>
<td>(8.146)</td>
<td>(9.044)</td>
</tr>
<tr>
<td>Total Sales</td>
<td>84.1</td>
<td>0.043</td>
<td>0.034</td>
</tr>
<tr>
<td>(million $)</td>
<td></td>
<td>(0.009)</td>
<td>(0.010)</td>
</tr>
<tr>
<td>Interest Cost</td>
<td>14.4</td>
<td>-1.376</td>
<td>-1.906</td>
</tr>
<tr>
<td>(percent)</td>
<td></td>
<td>(0.756)</td>
<td>(0.835)</td>
</tr>
<tr>
<td>Borrowed Reserves</td>
<td>1.5</td>
<td>5.587</td>
<td>8.704</td>
</tr>
<tr>
<td>(billion $)</td>
<td></td>
<td>(3.792)</td>
<td>(4.399)</td>
</tr>
<tr>
<td>Collateral</td>
<td>0.2</td>
<td>2.693</td>
<td>3.426</td>
</tr>
<tr>
<td></td>
<td></td>
<td>(2.875)</td>
<td>(3.312)</td>
</tr>
<tr>
<td>$R^2$</td>
<td></td>
<td>0.258</td>
<td>—</td>
</tr>
<tr>
<td>SEE/$a$</td>
<td></td>
<td>11.172</td>
<td>12.104</td>
</tr>
<tr>
<td>Sample Size</td>
<td></td>
<td>90</td>
<td>90</td>
</tr>
</tbody>
</table>

Note: The dependent variable is desired borrowing in million $. Standard errors for least squares estimates (asymptotic standard errors for MLE) appear in parentheses below the coefficients.

A The average value maximum take-down is $5.1 million

V. Conclusion

From our survey of loan contracts, we report on some important features of loan contracts. We then use our micro contract data to estimate a model of borrowing behavior. The parameter estimates are of the expected sign and are generally statistically significant.

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EXAMINING THE RELATIONSHIP BETWEEN CAPITAL INVESTMENT AND HOSPITAL OPERATING EXPENDITURES

Gerard Anderson, Jane Erickson, and Susan Feigenbaum*

Abstract—Recent hospital payment initiatives necessitate a precise estimate of the relationship between hospital capital expenditures and operating costs. Most previous estimates have ignored the possibility that capital and operating costs may be determined simultaneously and that the impact may occur over several years. We estimate several models and show that previous estimates are based upon a specific set of assumptions which should be examined carefully. Our preferred specification estimates a much lower effect of capital expenditures on operating costs than previous models.

The relationship between hospital capital and operating expenditures has become increasingly important in recent years for both policymakers and providers. Policymakers are questioning a long established premise of health planning (certificate of need) programs that controlling capital expenditures will ultimately result in lower operating costs. Congress is debating whether to incorporate capital into the Medicare prospective payment system. Hospital administrators, confronted for the first time with a price sensitive environment, are becoming increasingly concerned about the impact of capital projects on current and future operating expenditures.

Empirical studies suggest that capital expenditures increase both short- and long-run operating costs, although the estimated effect varies widely. This variation in the empirical estimates may be attributable to the fact that most previous studies have ignored the possibility that capital and operating outlays may be simultaneously determined and that the impact may occur over several years.

The purpose of this paper is to explore how these specification issues influence the estimated relationship between capital and operating expenditures. Section I reviews the findings of previous studies. Section II discusses theoretical considerations related to the development of a hospital operating cost function, focusing on the appropriate treatment of capital. Parameter estimates for alternative specifications of this function are presented in section III. In conclusion, we comment on the policy implications of our findings.

I. Prior Studies

Numerous researchers have generated estimates of the impact of capital investment on hospital operating expenditures. Somers (1969) estimated that for every dollar spent on capital, a hospital’s operating expenses would increase 35 to 40 cents in the same year, while Dunn and Lefkowitz (1978) estimated the increase to be 50 cents in the following year. More recently, Bentkover et al. (1984) estimated that for every dollar invested in capital, current operating costs increased by 22 cents and, over a ten year period, the net present value of operating costs increased by $1.84.

Some of the earlier work fails to consider factors other than capital investment that influence operating costs which, if omitted, may confound the estimated impact of capital outlays on operating expenditures. Hospital characteristics such as bed size, ownership, wage levels, case mix and teaching status, as well as the presence of such regulatory programs as state rate-setting and certificate of need, can impact both capital investment and operating expenditures.1

Most previous studies also have ignored the potentially simultaneous relationship between capital and operating costs (Pauly, 1974). Recognizing this possible interdependence, Furst and Markland (1980) incorporated a lead/lag structure in their analysis to distinguish whether (i) operating cost increases follow capital investment, reflecting the creation of demand for hospital services; or (ii) operating costs rise prior to new capital outlays, signalling full utilization of existing capacity and a need for additional capital. Their em-