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BORROWING CONSTRAINTS AND FEMALE LABOR SUPPLY : NONPARAMETRIC AND PARAMETRIC EVIDENCE OF THE IMPACT OF MORTGAGE LENDING RULES

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RÉSUMÉ

Dans cet article, on présente des résultats empiriques qui démontrent que les exigences de revenu relatives aux prêts hypothécaires induisent des distorsions significatives sur l'offre de travail des femmes. Les politiques économiques affectant les taux d'intérêt, de même que les politiques fiscales facilitant la mise de fonds, ont une influence directe sur cette contrainte à l'emprunt. On présente d'abord des statistiques mettant en rapport les choix immobiliers des ménages et la participation des femmes au marché du travail. Ces statistiques indiquent que les choix des ménages ne correspondent pas aux prédictions du modèle traditionnel des choix de consommation et de loisir sur le cycle de vie. En second lieu, des régressions non paramétriques de type "kernel" comparant l'offre de travail des propriétaires et des locataires illustrent les distorsions induites par les exigences de revenu relatives aux prêts hypothécaires. Enfin, on évalue l'impact du rapport d'amortissement brut de la dette et de la contrainte à l'emprunt sur l'offre de travail des femmes à l'aide d'un modèle en forme réduite. L'endogénéité de ces variables est testée et rejetée en faisant appel à une procédure en deux étapes, dérivée d'une approche de maximum de vraisemblance conditionnel, qui utilise des résidus généralisés pour effectuer un test de spécification à la Hausman. Des estimés asymptotiquement valides des écarts types sont obtenus à l'aide d'une procédure bootstrap non paramétrique.

Mots clés : contraintes de liquidité, offre de travail des femmes, régression kernel, résidus généralisés, bootstrap.

ABSTRACT

This paper presents empirical evidence of significant distortions induced by the earnings test of the mortgage qualification process on the labor supply of married women. This borrowing constraint is directly affected by changes in interest rates and by fiscal rules facilitating the provision of a down payment. First, basic statistics on housing choices and female labor market participation are shown to be inconsistent with the traditional life-cycle model of consumption-leisure choices. Second, kernel regressions, contrasting the labor supply of the wives of homeowners and of renters, are used to illustrate the distortion brought about by the earnings test. Finally, the impact of the debt-service ratio and of the borrowing constraint on the wives' labor supply is assessed using a reduced-form approach in the spirit of so-called second generation empirical studies of female labor supply. The endogeneity of these variables is tested and rejected using a two-stage procedure, derived from a conditional maximum likelihood approach, that relies on generalized residuals to perform a Hausman-type specification test. Nonparametric bootstrapping is used to obtain asymptotically valid estimates of the standard errors. Implications of the results for the analysis of life-cycle behavior and housing choices are discussed.

Key words : borrowing constraints, female labor supply, kernel regression, generalized residuals, bootstrapping.

I. Introduction

The purpose of this paper is to present empirical evidence of significant distortions induced by borrowing constraints on the labor supply of married women. In recent years, empirical attempts to explain the excess sensitivity —with reference to the life-cycle/permanent income hypothesis— of consumption to current income by the presence of liquidity constraints have yielded mixed results. Using micro data, Flavin (1985) and Zeldes (1989) among others, found some evidence of liquidity constraints, while Altonji and Siow (1987), Runkle (1991), and Flavin (1991) found no evidence that liquidity constraints lead to the rejection of the permanent income hypothesis. Related inquiries into the effects of liquidity constraints on the intertemporal labor supply response of primary male earners (Ball, 1990, Dau-Schmidt, 1992) have yielded equally meager results.

Yet, it is not surprising that such results are elusive. In these studies, liquidity constrained households are defined more or less arbitrarily as households holding low levels of liquid assets. But it is not known what levels of liquid assets are sufficient to allow consumption smoothing. [2] Focusing on housing assets may prove to be a more revealing avenue of research. In many countries, housing assets are held by a great majority of households for whom they are their dominant asset. [3] And more importantly, the borrowing constraints imposed on prospective homebuyers by mortgage lending institutions are well-defined.

In fact, Engelhardt (1991) found that downpayment constraints on homes induced significant distortions in the consumption of homebuyers. Dau-Schmidt (1992) found that primary male workers with high consumption commitments, including housing, car, and other debt payments, had substantially lower in-

[2] As a result, liquidity constrained households may have been included in the group assumed not to be constrained. In Runkle (1991), homeowners are assumed not to be liquidity constrained. However, reducing home equity and thereby increasing housing payments may not be possible for some homeowners who need to renew their mortgage. In the 1986 Canadian FAMily EXpenditures survey, households who borrowed against their home equity borrowed on average $13600 or 27% of their total expenditures for the year.

[3] For the U.S., see Skinner (1991).

tertemporal substitution elasticities than non-constrained workers. In Fortin (1992), I showed that mortgage commitments induced a significant portion of married women to participate in the labor market. In that study, I also estimated a life-cycle consistent model of household labor supply, that incorporated a mortgage qualification constraint based on earnings, among households with a participating female. I found that the households' housing choices were reliant on the existing levels on the wife's labor earnings.

It is difficult to distinguish workers who, because of high levels of anticipated labor supply, make high commitments, from workers who, because of their high commitments, are bound to equally high level of labor supply. Yet, this distinction is crucial if the potentially strong work incentive effects of mortgage commitments are to be incorporated into public policies and programs. This paper seeks to establish that, as households' mortgage choices are limited by borrowing constraints, the labor supply of married women is distorted by the mortgage qualification constraint based on earnings. This constraint is directly affected by changes in interest rates and by fiscal rules facilitating the provision of a downpayment.

The empirical analysis utilizes the 1986 Canadian Family Expenditures Survey (FAMEX). There are many advantages to using a Canadian data base: the absence of a preferential tax-treatment of owner-occupied housing, the relative uniformity of mortgage lending rules, the prevalence of very short term adjustable rate mortgages that make the effect of mortgage lending rules more pervasive. [4] However, an analysis from a cross-section of cohorts will tend to understate the effects of mortgage lending rules on the households that are directly affected by policies encouraging homeownership. [5]

I present the argument as follows. In section 2, I exploit regional differences in patterns of homeownership to present evidence of the credit crunch on home-

[4] The mortgage lending rules are largely set by the principal mortgage insurer, the Canada Central Mortgage and Housing Corporation.

[5] Panel data is unfortunately not available for Canada.

owners. Double income households sampled in Western Canada live in homes of lesser value on average than households with a nonparticipating female. In section 3.1, kernel regressions, contrasting the labor supply of the wives of homeowners and of renters, are used to illustrate the distortion brought about by the mortgage qualification constraint. In section 3.2, the issue of endogeneity of mortgage choices is addressed via a classic instrumental variables treatment of the wives' labor supply. To test the endogeneity of the debt-service ratio and of the borrowing constraint on the wives' labor supply, I use a two-stage procedure, derived from the conditional maximum likelihood approach of Smith and Blundell (1986), that relies on generalized residuals (Gouriéroux *et al.* 1987, Pagan and Vella, 1989) to perform a specification test in the spirit of Hausman (1978). Nonparametric bootstrapping is used to obtain asympotically valid estimates of the standard errors. Both the standard Nelson and Olsen (1978) estimates and the ones from the modified Smith and Blundell procedure strongly support the weak exogeneity of the borrowing constraint. Finally, in section 4, I comment on the implications of the results.

II. Housing Demand, Borrowing Constraints, and Married Women Labor Market Participation

1. Analytical Framework

In the standard life-cycle model, the household's labor market and housing consumption decisions result from an optimization process that links these decisions to the household's lifetime wealth. Housing investment decisions are made according to an optimal portfolio rule and do not necessarily coincide with housing consumption decisions. The implausibility of such a simple framework to analyze housing decisions has prompted many authors to introduce more realistic factors. They have been concerned with the distorting role of financing imperfections and with peculiarities of the housing market. The downpayment constraint has been shown to force households with high lifetime wealth but low current liquid wealth to make non optimal housing choices (Jones, 1990). The equality of housing investment and housing consumption is a complication that has been shown to cause portfolio imbalances (Henderson and Ioannides, 1983), when housing levels are primarily determined by consumption demand for housing. In a general model of optimal consumption and portfolio selection with an illiquid durable good such as housing, Grossman and Laroque (1990) have shown that small moving costs can introduce an element of considerable rigidity into housing decisions. Kearl (1978) has studied the impact on the demand for housing assets of the "tilting" of real mortgage payments, that results from the combination of fixed nominal mortgage interest and high inflation rates.

Here, I choose to focus on the potential distortions induced on labor supply by the earnings test of the mortgage qualification process. In most countries, mortgage lending institutions require that housing payments not exceed a proportion, usually around one third, of the household's income. In the presence of high housing prices, this earnings test may supersede the downpayment constraint as it may force prospective homebuyers to come up with a downpayment larger that the minimum required. Alternatively, it may force married women to return temporarily to work in order to afford a bigger mortgage. It is not difficult to find anecdotal evidence of such behavior. In Canada, mortgage contracts are of the adjustable rate variety and need to be renegotiated every 2-5 years, thus this mortgage lending rule is expected to have important behavioral effects. In the U.S., because the share of adjustable rate mortgages of all mortgages has varied considerably in the 1980's (from a high of 60% in 1986 to a low of 20% in 1987), the following findings may apply with less strength. Similarly, the impact of the earnings test on the labor supply of secondary earners may be less important in the U.K, where many lenders give the income of the bigger earner greater weight, sometimes three times the weight (Smith, 1991).

The earnings test introduces an additional constraint into the traditional life-cycle model of household's consumption–leisure choices. From the time ($t = 0$) the house is purchased, if the mortgage needs to be renewed periodically, the

household's income cannot fall below a maximum allowable proportion $1/k$ of the mortgage payments, denoted for expositional simplicity by a constant annuity M_0. [6] Then, under complete certainty, the homeowner's problem can be written in the optimal control format, where constraints are easily incorporated, as in (Fortin, 1992):

$$\begin{aligned}
\text{Maximize}_{\{l_{mt}, l_{ft}, q_t\}} \quad & \sum_{t=1}^{L} \delta^{(t-1)} U_t(l_{mt}, l_{ft}, q_t) \\
\text{subject to} \quad & W_{t+1} - W_t = -\rho_t(w_{mt} l_{mt} + w_{ft} l_{ft} + p_t q_t) \\
& k[w_{mt}(T_m - l_{mt}) + w_{ft}(T_f - l_{ft})] - M_0 \geq 0 \\
& W_1, W_L \quad \text{fixed.}
\end{aligned} \tag{2.1}$$

where δ is a subjective rate of discount, where ρ_t is the market discount rate, where l_{mt} and l_{ft} represent male and female leisure, respectively, and q_t represents non-housing consumption, with respective prices w_{mt}, w_{ft}, and p_t, where constant time endowments T_m and T_f are assumed for all periods, where W_t is the usual discounted present value of human and nonhuman wealth, exclusive of housing wealth, and where preferences are usually parameterized to accommodate demographic characteristics. [7] In this theoretical model, the household's optimal leisure (or hours of work) choices depend not only on the household's marginal utility of wealth, on the spouses' respective wages, on the other spouse's earnings, but also, when the mortgage qualification constraint is binding, on their mortgage payments. In practice, the household's labor market choices are expected to be gradually constrained as the ratio of mortgage payments to family income approaches the allowable limit.

Conversely in a model with endogenous tenure choice, the household's housing choices will be partially determined by its labor market choices. According

[6] Before the home is purchased, the downpayment constraint may introduce some distortions in the household's labor market decisions. However, the analysis of such distortions would require not only panel data, but information on renters' plans to purchase a home, as in Yoshikawa and Ohtake (1989).

[7] Housing consumption is derived from the housing choice made at the time $t = 0$ and is thus separable from other consumption-leisure choices. There also might be inequality constraints on the choice variables.

to Henderson and Ioannides' (1987) model of owner-occupancy, if the investment motive dominates housing decisions, one would expect the household to hold housing assets in amounts exceeding its consumption demand. This would apply with more strength where housing assets carry important tax-advantages in comparison to other assets. Households with high anticipated levels of lifetime wealth would choose high levels of housing assets, which could be seen as endogenous in their labor market decisions. [8] On the other hand, if the household's consumption demand for housing services exceeds its asset demand, owner occupancy would force its housing assets up to the level of its consumption demand. In this case, the lumpiness of housing assets together with borrowing constraints could distort the household's labor market choices. The household's labor market decisions would thus be constrained by its housing commitments, which could then be treated as exogenous as in the model (2.1) above.

I first present some simple facts on the interrelation between housing choices and the labor market participation of wives. Since the latter is more variable than that of husbands, it may capture better the household's labor market adjustments to varying circumstances. These facts show choices that are inconsistent with predictions from the traditional life-cycle model and that may be the result of borrowing constraints.

2. Data

The empirical evidence is drawn from the 1986 FAMEX survey, which contains detailed information describing the family unit, its housing arrangements, various financial variables, the labor status of its members, and a somewhat detailed account of expenditures per commodity categories. There are, however, limitations imposed by FAMEX data base. While it is possible to identify the movers-purchasers, it is not known at what date nonmovers purchased their home. Second, the information on household location is provided very crudely for only six regions. [9]

From an original sample of 10,356 households, 3,721 households were removed because they lived in agglomerations of less than 100,000 inhabitants, or because the area they lived in was masked. Since the focus of the study is on women as non-primary earners, 2,542 households were excluded because they were made of unattached individuals, single parent families, or other non-family units. Another 556 households were deleted because the husband was retired. Of the remaining 3537 households, 2562 households were homeowners.

TABLE I
MARRIED WOMEN LABOR FORCE PARTICIPATION RATES BY AGE, FAMILY TENURE AND REGION

Age	Owners without a Mortgage		Owners with a Mortgage		Tenants		Changed from[a] Tenant to Owner	
	#	%	#	%	#	%	#	%
				East[b]				
≤ 24	2/2	100.0	21/26	80.8	111/144	77.1	15/18	83.3
25-34	57/78	73.1	300/394	76.1	205/273	75.1	41/53	77.4
35-44	109/170	64.1	279/370	75.4	71/105	67.6	10/15	66.7
45-64	158/302	52.3	148/237	62.5	55/105	52.4	8/10	80.0
All	326/552	59.1	748/1027	72.8	442/627	70.5	74/96	77.1
				West[c]				
≤ 24	1/1	100.0	16/19	84.2	61/73	83.6	16/16	100.0
25-34	26/37	70.3	144/183	78.7	113/141	80.1	20/22	90.9
35-44	49/70	70.0	173/211	82.0	34/49	69.4	9/10	90.0
45-64	86/157	54.8	97/140	69.3	22/33	66.7	3/5	60.0
All	103/265	61.1	430/554	77.8	230/296	77.7	48/53	90.6

Note: This table exclude 21 households with wives aged 65 and up.

[a] Includes families who changed tenure, from tenant to homeowner, in 1986.

[b] East includes Atlantic Provinces, Québec and Ontario.

[c] West includes Manitoba, Sasketchewan, Alberta and British Columbia.

[8] Philips and Vanderhoff (1991) provide empirical evidence from the PSID supporting that view for two-earner households in which the wife is employed in a professional or managerial occupation. These households spend more on housing for given levels of total family income.

[9] Further, that information is masked for a few (6) households with high home values (> $250 000)

3. *Basic Facts*

Table I shows that the labor market participation of married women is about 10 points higher, for women aged 35 up, among homeowners with a mortgage than among homeowners without a mortgage and among families who are tenants. [10] This result holds both in the East, defined as all provinces east of and including Ontario, and in the West, defined as all provinces west of and including Manitoba. [11] In Fortin (1992), I also showed using a Probit model of labor market participation that this result holds even while correcting for other variables, such as the value of the home owned, the husband's income, the number of children, etc. Further, I showed that the ratio of mortgage payments to family income, exclusive of the wive's labor income, had a strong nonlinear effect on the wives' labor market participation as the ratio approached the allowable limit.

Table II shows there are significant differences between families with a participating female and families with a nonparticipating female in the means of various housing and financial variables. In the East, families with a participating female have significantly higher incomes on average and live in houses of significantly higher values on average than families with a nonparticipating female. In the West, families with a participating female have significantly higher incomes on average than their counterparts, but they live in houses of lesser values on average. Though the level of significance of this latter result is not high, it contradicts the life-cycle implication that households with a participating wife buy the bigger house that their higher lifetime wealth allows them

[10] A woman is a market participant if the number of weeks worked full time or part time is greater than zero and if her labor income is greater than zero.

[11] This geographical division is used in section II to get enough observations per cell in Table II and III. However, the broad results generally hold when the division is done along the six regional categories available in the survey. These regional categories will be reintroduced in the more rigorous analysis of section 3.

TABLE II
MEANS OF VARIABLES BY REGION

	Owners without a Mortgage			Owners with a Mortgage		
	Nonparticipating Female	Participating Female	F-value	Nonparticipating Female	Participating Female	F-value
Value of Home Owned						
East[a]	102 291 (75 589)	123 247 (76 775)	10.20**	95 972 (57 167)	105 731 (59 682)	6.08**
West[b]	114 323 (66 671)	97 932 (54 140)	4.95**	101 239 (52 296)	94 023 (42 375)	2.68*
Balance of Principal on Mortgage						
East				29 453 (21 792)	36 516 (24 532)	19.55**
West				38 095 (28 070)	42 389 (26 151)	2.68*
Family after Taxes Income						
East	34 855 (17 942)	48 096 (20 506)	62.43**	34 241 (14 919)	42 604 (14 969)	69.55**
West	36 393 (17 041)	45 004 (18 517)	15.01**	35 565 (16 688)	42 540 (14 932)	21.28**
Family Investment Income						
East	3 266 (8 279)	3 067 (6 091)	0.11	1 605 (8 086)	1 031 (3 902)	2.55
West	5 649 (9 292)	3 942 (7 994)	2.60*	918 (18 44)	1 169 (4 267)	0.43
Number of Rooms						
East	6.67	7.26	14.94**	7.06	6.94	1.14
West	7.37	7.12	1.06	7.53	7.33	1.19
Wife's Age						
East	48.8	44.1		40.1	36.7	
West	50.3	45.7		40.9	36.9	
Number of Observations						
East	231	326		305	822	
West	109	162		131	476	

Note: Standard deviations in parentheses. Analysis of variance indicates that the differences between households with and without participating female are statistically significant at the 5% level (**) and the 10% level (*).

[a] East includes Atlantic Provinces, Québec and Ontario.

[b] West includes Manitoba, Sasketchewan, Alberta and British Columbia.

to. It is not possible to exclude the simple explanation of possible difference in tastes among eastern and western households, especially in terms of portfolio diversity. However, the existence of borrowing constraints is another daunting explanation.

Table III focuses on movers and new homeowners to investigate this puzzle. There are differences between the East and the West in the ratio of the downpayment to the value of the house. This ratio is significantly lower in the West among all movers, especially among movers and new owners with a participating female, and cannot be explained in this sample by houses of higher value in the West. Rather, as the public transit systems are less efficient in the West, the need for a second car in households with a participating female could explain why new homebuyers were unable to accumulate a downpayment as large as their eastern counterparts. In fact, the median downpayment among new western homebuyers is at the 25% of home value limit below which costly mortgage insurance is required. These results are in line with those of Jones (1990) who concluded that, in Canada, current liquid wealth was a more important determinant of housing demand among young owners than labor earnings.

These basic statistics indicate that the endogeneity of housing choices in labor market decisions, implied by the unconstrained life-cycle model, cannot be held as straightforward assumption. This issue will be investigated more thoroughly in section 3.2. The next subsection simply seeks to describe the impact of the earnings test on the labor supply on married women.

TABLE III
MEANS OF VARIABLES AMONG HOMEBUYERS BY REGION

	Movers			New Owners		
	Nonparticipating Female	Participating Female	All	Nonparticipating Female	Participating Female	All
Value of Home Owned						
East	109 417	125 452	120 251	70 900	96 506	90 687
	(73 257)	(56 336)	(62 276)	(39 514)	(37 823)	(39 486)
West	119 727	116 984	117 686	118 300	82 770	86 253
	(67 642)	(51 220)	(55 020)	(74 518)	(35 032)	(40 775)
F-value	0.16	0.47	0.05	3.98*	3.84*	0.40
Downpayment as a Proportion of the Value of Home Owned						
East[a]	.61 (.29)	.57 (.26)	.59 (.27)	.27 (.19)	.37 (.19)	.35 (.19)
West[b]	.47 (.29)	.42 (.31)	.44 (.30)	.41 (.19)	.29 (.18)	.30 (.18)
F-value	1.64	5.82**	7.85**	2.13	5.86**	2.31
Number of Cars						
East	1.3 (.69)	1.8 (.68)	1.6 (.72)	1.6 (.82)	1.3 (.67)	1.4 (.72)
West	1.5 (.69)	1.8 (.71)	1.7 (.70)	1.2 (.45)	1.8 (.69)	1.7 (.69)
F-value	1.02	0.00	0.53	1.08	14.58**	8.87**
Number of Observations						
East	24	50	74	20	68	88
West	11	32	43	5	46	51

Note: Standard deviations in parentheses. Analysis of variance indicates that differences between East and West are statistically significant at the 5% level (**) and at the 10% level (*).

[a] East includes Atlantic Provinces, Québec and Ontario.

[b] West includes Manitoba, Saskatchewan, Alberta and British Columbia.

III. The Impact of Mortgage Lending Rules on the Labor Supply of Married Women

1. *Nonparametric Evidence*

If the earnings test has a distorting effect on the labor supply of married women, wives should increase their labor supply as the ratio of housing payments to family income, exclusive of their labor earnings, approaches the allowable limit. I use kernel regressions of the number of weeks worked by wives (denoted Y) as a function of this ratio (denoted X) to describe the potential nonlinearities induced by the borrowing constraint. Kernel regressions are obtained by computing what amounts to be moving averages, using kernel weights, of the number of weeks worked for each value of the ratio X. Among other smoothing techniques, kernel estimators are often preferred by econometricians because their asymptotic properties are fairly well established (Bierens, 1987) and because they seem to have good finite sample properties. [12]

More formally, given an identically independently distributed data set $\{(X_i, Y_i)\}_{i=1}^n$, kernel smoothing approximates the mean response curve m of the relationship

$$Y_i = m(X_i) + \epsilon_i, \quad i = 1, \ldots, n \tag{3.2}$$

where $E(\epsilon|X) = 0$, using the Nadaraya-Watson estimator

$$\hat{m}_h(x) = n^{-1} \sum_{i=1}^{n} W_i(x) Y_i = \frac{n^{-1} \sum_{i=1}^{n} K_h(x - X_i) Y_i}{n^{-1} \sum_{i=1}^{n} K_h(x - X_i)} \tag{3.3}$$

where the weight functions for kernel smoothers are given by

$$W_i(x) = K_h(x - X_i)/\hat{f}_h(x), \tag{3.4}$$

[12] Moreover, other smoothing techniques, such as cubic splines, have been shown to be asymptotically equivalent to kernel smoothers (Silverman, 1986).

with kernel density estimate and kernel function given by

$$\hat{f}_h(x) = n^{-1} \sum_{i=1}^{n} K_h(x - X_i) \quad \text{and} \quad K_h(u) = h^{-1} K(u/h) \tag{3.5}$$

where the scale factor h is called the bandwidth. The kernel function $K(\cdot)$ is chosen to be a real continuous function which is bounded, symmetric around zero, and integrates to one.

An often heard critic of kernel estimation is that the shape of the kernel regression is very sensitive to the choice of kernel function and bandwidth. The choice of kernel function is less problematic than the choice of bandwidth. Among the different kernel functions of order 2, the Epanechnikov kernel,

$$K(u) = 0.75(1 - u^2) I(|u| \leq 1). \tag{3.6}$$

which is of parabolic shape, has been shown (Silverman, 1986, Härdle, 1990) to be the most efficient in the sense of minimizing the approximate mean integrated square error.

The choice of bandwidth determines the optimality of the kernel estimator. When h is too small, the resulting curve is too wiggly $[\hat{m}_h(X_i) \to Y_i$ as $h \to 0]$. When h is too large, the resulting curve may smooth away some important features of the data $[\hat{m}_h(x) \to \bar{Y}$ as $h \to \infty]$. While the development of automatic bandwidth selectors remains a topic of intense research (Park and Turlach, 1992), two quadratic measures of accuracy have been widely used. Choosing h that minimizes the Cross-Validation function

$$CV(h) = n^{-1} \sum_{j=1}^{n} (Y_j - \hat{m}_{h,j}(x))^2, \tag{3.7}$$

where $\hat{m}_{h,j}(x) = n^{-1} \sum_{i \neq j} W_i(x) Y_i$, has been shown to be asymptotically optimal. [13] This leave-one-out method avoids the problem of using Y_i twice in

[13] In the sense that the ratio of estimated loss to minimum loss asymptotically tends to one $d_A(h^*)/inf_h\, d_A(h) \xrightarrow{a.s.} 1$. See, Härdle (1990).

approximating the average square error

$$d_A(h) = n^{-1} \sum_{j=1}^{n} (\hat{m}_h(X_j) - m(X_j))^2. \tag{3.8}$$

The CV function also may be weighted to down play boundary effects.

Another widely used bandwidth selector is the Generalized Cross-Validation function

$$\begin{aligned} GCV(h) &= \Xi_{GCV}(n^{-1}h^{-1})\ p(h) \\ &= (1 - n^{-1}h^{-1}K(0))^{-2}\ n^{-1} \sum_{j=1}^{n} (Y_j - \hat{m}_j(x))^2, \end{aligned} \tag{3.9}$$

which uses a penalizing function $\Xi_{GCV} = (1 - n^{-1}h^{-1}K(0))^{-2}$ to correct the biasedness of the prediction error $p(h)$ as an estimator of $d_A(h)$. While bandwidth selectors are all asymptotically optimal, their speed of convergence varies.

Figure 1 displays kernel regression estimates of the number of weeks worked by married women (Y) as a function of the ratio (X) of housing payments (or rent payments) to family income, exclusive of the woman's labor income, for choices of bandwidth that minimize the CV function. The kernel regressions are shown for all homeowners, for homeowners with a mortgage and for tenants. As expected, the labor supply of married women responds positively to increases in X. [14] The more striking feature, among homeowners, is the singular increase in slope beginning at $X = 0.2$ that brings the wife's number of weeks worked to a higher plateau as X exceeds 0.25. [15] This pattern could correspond to the

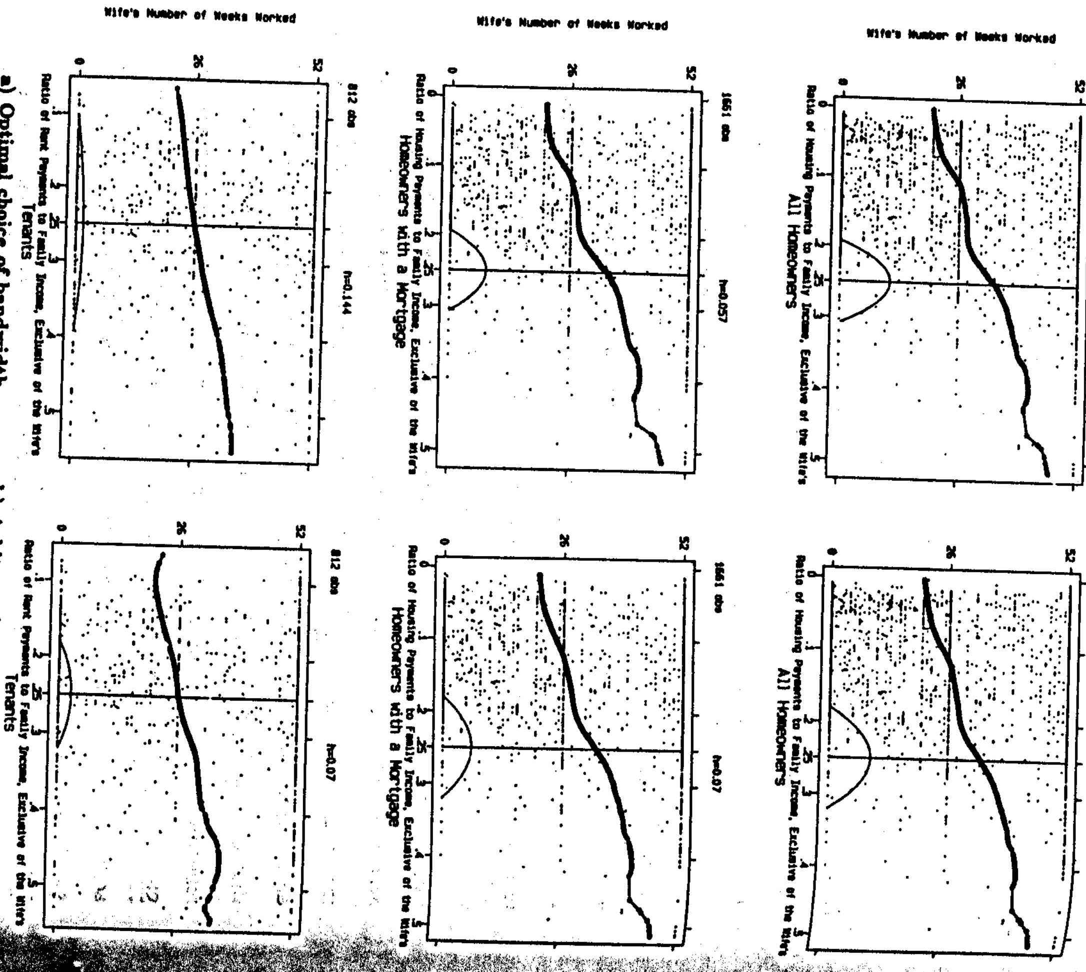
[14] These figures leave out observations for which the ratio X was greater than 0.5, for homeowners, and 0.55, for tenants, since the density of the observations was becoming too low. The observations deleted include households in which the wife was the sole earner in 1986. The calculations were made with a program written in STATA by the author.

[15] The FAMEX survey gives the number of weeks worked full-time and the number of weeks worked part-time. To obtain a measure of the total number of weeks worked, the number of weeks worked part-time were given half the weight of the number of weeks worked full-time while

empirical manifestation of a rationing constraint. The multiple-plateau shape of the kernel regression function among homeowners is in sharp contrast with the smoothly increasing kernel regression function among tenants.

Figure 1 also shows the effective kernel weights used at $X = 0.25$. These effective kernel weights may help the reader assess the "reasonableness" of the bandwidth selection. Table IV reports the values of the cross-validation function $CV(h)$, of the generalized cross-validation function $GCV(h)$, and of a more common quadratic measure of accuracy, the mean square error, $MSE(h) = (n-DF)^{-1}\sum_{j=1}^{n}(Y_j - \hat{m}_j(x))^2$, where the degrees of freedom are $DF = K(0)/h$ (as in Zheng, 1991). Since the MSE is a biased measure of $d_A(h)$, it should not be used as a guide in the choice of the optimal bandwidth. The window width that ensure the maximum rate of convergence of the kernel estimator to its asymptotic normal distribution, is given by $h = cn^{-1/5}$ (Bierens, 1987), where the optimal constant c is the one balancing the unknown squared bias and variance. For practical purpose, the grid search for the optimal value of h is done over values equal to $i \cdot 0.05 \cdot n^{-1/5}$, $i = 1, 2, \ldots$. The bandwidths that minimize the cross-validation function were $h = 0.058$ among all homeowners, $h = 0.0567$ among homeowners with a mortgage, $h = 0.144$ among tenants. The bandwidths that minimize the generalized cross-validation function were somewhat higher, reflecting the fact the measures of accuracy are only asymptotically equivalent. Since the optimal bandwidth for tenants was much larger than for the groups, additional kernel regressions were run with the same window width ($h = 0.07$) for all data sets.

adding the two measures up to a maximum of 52 weeks. This explain why there are a large number of observations at 26 weeks. This weighting scheme explains why the second plateau is at a little over 26 weeks in Figure 1. While alternative weighting schemes would change the level of the lower plateau, thereby attenuating or accentuating the incline between $X = 0.2$ and $X = 0.25$, they do not change the overall shape of the kernel regression function.



a) Optimal choice of bandwidth b) Arbitrary bandwidth (h=0.07)

Figure 1. Kernel estimate of the regression of the ratio of housing payments to family income, exclusive of the wife's labor earnings on the number of weeks worked by the wife (solid line), with effective kernel weight $K_h(x-\cdot)/\hat{f}_h(x)$ (lighter line) at $x = 0.25$ for an optimal choice of h (a) and for an arbitrary h (b) with Epanechnikov kernel $K(u) = 0.75(1-u^2)I(|u| \leq 1)$.

TABLE IV

QUADRATIC ERROR MEASURES FOR BANDWIDTH (h) SELECTION IN THE KERNEL REGRESSION OF THE NUMBER OF WEEKS WORKED BY MARRIED WOMEN AS A FUNCTION OF THE RATIO OF HOUSING PAYMENTS TO FAMILY INCOME, EXCLUSIVE OF THE WIFE'S LABOR EARNINGS

All Homeowners					Homeowners with a Mortgage					Tenants				
h	DF	MSE[a]	CV[b]	GCV[c]	h	DF	MSE	CV	GCV	h	DF	MSE	CV	GCV
0.0053	142.24	461.4787	461.6611	521.5664	0.0057	132.19	448.6888	449.2170	529.6352	0.0655	11.46	439.3206	439.3563	451.98
0.0105	71.12	454.5389	453.6437	482.7735	0.0113	66.09	439.6401	438.8841	476.8334	0.0786	9.55	439.2463	439.2486	449.75
0.0158	47.41	452.5297	451.6739	470.9854	0.0170	44.06	436.7994	436.2034	460.9302	0.0851	8.81	439.1564	439.1467	448.84
0.0211	35.56	451.0416	450.1825	464.7338	0.0227	33.05	435.0957	434.3006	452.9398	0.0917	8.18	439.1406	439.1212	448.12
0.0264	28.45	450.4295	449.7088	461.3190	0.0284	26.44	434.0362	433.3179	448.1901	0.0982	7.64	439.2383	439.2108	447.61
0.0316	23.71	449.7431	449.1298	458.7767	0.0340	22.03	433.3158	432.7138	445.0436	0.1047	7.16	439.3039	439.2704	447.15
0.0369	20.32	449.4165	448.9027	457.1374	0.0397	18.88	433.1468	432.6026	443.1664	0.1113	6.74	439.2131	439.1772	446.59
0.0422	17.78	449.3883	448.9293	456.1328	0.0454	16.52	433.0968	432.5800	441.8440	0.1178	6.36	439.0820	439.0473	446.04
0.0475	15.80	449.3153	448.9038	455.3019	0.0511	14.69	432.8421	432.3622	440.5998	0.1244	6.03	439.0070	438.9756	445.60
0.0527	14.22	449.2070	448.8358	454.5883	0.0567	13.22	432.6783	432.2289	439.6482	0.1309	5.73	438.9164	438.8898	445.17
0.0580	12.93	449.1213	448.7863	454.0085	0.0624	12.02	432.6565	432.2533	438.9855	0.1375	5.46	438.8367	438.8161	444.79
0.0633	11.85	449.0921	448.8004	453.5687	0.0681	11.02	432.6709	432.3216	438.4674	0.1440	5.21	438.8070	438.7937	444.48
0.0685	10.94	449.0863	448.8365	453.2161	0.0738	10.17	432.8015	432.4989	438.1496	0.1506	4.98	438.8318	438.8274	444.26
0.0738	10.16	449.0990	448.8848	452.9320	0.0794	9.44	433.0034	432.7437	437.9686	0.1571	4.77	438.8988	438.9038	444.10
0.0791	9.48	449.1765	448.9956	452.7532	0.0851	8.81	433.2745	433.0582	437.9090	0.1637	4.58	438.9612	438.9753	443.95
0.0844	8.89	449.3194	449.1707	452.6723	0.0908	8.26	433.5825	433.4051	437.9282	0.1702	4.41	439.0865	439.1092	443.89
0.0896	8.37	449.4924	449.3725	452.6483	0.0965	7.78	433.8867	433.7393	437.9778	0.1768	4.24	439.2459	439.2766	443.87
0.0949	7.90	449.6938	449.5970	452.6748	0.1021	7.34	434.1845	434.0598	438.0494	0.1833	4.09	439.4243	439.4630	443.88
0.1002	7.49	449.9315	449.8527	452.7564	0.1078	6.96	434.5237	434.4143	438.1869	0.1899	3.95	439.6470	439.6934	443.95
0.1055	7.11	450.2098	450.1441	452.8945	0.1135	6.61	434.8534	434.7533	438.3349	0.1964	3.82	439.8828	439.9359	444.04

[a] $MSE(h) = (n - DF)^{-1} \sum_{j=1}^{n} \left(Y_j - \hat{m}_j(x)\right)^2$, where $\hat{m}_j(x)$ is the kernel estimate and DF the degrees of freedom.

[b] $CV(h) = n^{-1} \sum_{j=1}^{n} \left(Y_j - \hat{m}_{h,j}(x)\right)^2$, where $\hat{m}_{h,j}(x) = n^{-1} \sum_{i \neq j} W_{ni}(x) Y_i$ and where $W_{ni}(x)$ is the kernel weight.

[c] $GCV(h) = \left(1 - n^{-1} h^{-1} K(0)\right)^{-2} n^{-1} \sum_{j=1}^{n} \left(Y_j - \hat{m}_j(x)\right)^2$

TABLE V

DAVIDSON-MACKINNON TEST OF THE LINEAR VS. NON-PARAMETRIC SPECIFICATIONS OF THE WIVES' LABOR SUPPLY FUNCTION

	h	$\hat{\alpha}$	t-statistic	P-value
All Homeowners				
optimal bandwidth	0.0580	1.1181	2.137	0.033
oversmoothed	0.07	0.8723	1.577	0.115
Homeowners with a Mortgage				
optimal bandwidth	0.0567	1.5651	2.814	0.005
oversmoothed	0.07	1.3051	2.111	0.035
Tenants				
optimal bandwidth	0.144	-0.0459	0.068	0.945
undersmoothed	0.07	1.6709	1.644	0.101

The crucial difference in the kernel regression function between tenants and homeowners is the somewhat sharp increase in slope between $X = 0.2$ and $X = 0.25$ in the homeowners regression. In an attempt to link this bent to the earnings test of the mortgage qualification process, it is important to see whether the perceived nonlinearity is significantly different from the linear alternative. I use a standard Davidson-MacKinnon procedure (1981) to test the null hypothesis:

$$H_0 : Y_i = g(X_i) = a + bX_i + \epsilon_{0i} \quad \text{against} \quad H_1 : Y_i = m_h(X_i) + \epsilon_{1i}. \tag{3.10}$$

Since H_0 is linear, the t-statistic on $\hat{\alpha}$ in the artificial regression

$$Y_i - \hat{g}(X_i) = \alpha[\hat{m}_{\hat{h}}(X_i) - \hat{g}(X_i)] + \epsilon_i, \tag{3.11}$$

where $\hat{m}_{\hat{h}}$ is the optimal kernel estimate, provide a test of whether H_0 is true. Wooldridge (1992) has shown, in the case of a linear null hypothesis, the consistency of a Davidson-MacKinnon type test against a sequence of nonnested alternatives growing slowly enough toward the null.

The t-statistics given in table V confirm that, at the optimal bandwidth, the wives' labor supply function is nonlinear in the debt-service ratio X for homeowners, while it is linear for tenants. At the alternative bandwidth $h = 0.07$, the nonlinearity is still significant for homeowners with a mortgage. A more precise test of the statistical significance of the borrowing constraint focuses on the perceived threshold at $X = 0.25$. It is presented next.

2. *Parametric Evidence*

In this section, I follow a reduced-form approach in the spirit of so-called (Killingsworth, 1983) second generation empirical studies of female labor supply, which consider the husband's behavior as exogenous. In Fortin (1992), I pursued a more structural analysis and estimated a life-cycle consistent model of commodity demand and household labor supply under a mortgage qualification constraint. As shown in Mroz (1987), the critical issues in the estimation of female labor supply are the use of a censored econometric model and the exogeneity of the wife's wage rate. Because the wage rate of participating females is not directly available from the data and because I include households with nonparticipating females, I adopt a reduced form approach to the wage rate specification. That is, I replace the wage rate in the labor supply equation by its determinants: the wife's education, age, the number of children, etc. [16]

The preceding descriptive discussion revealed the existence of a threshold at $X = 0.25$ beyond which the wives' labor supply appeared to reach higher levels. This threshold potentially captures the institutional borrowing constraint. A

[16] In Fortin (1992), I used an alternative approach and obtained a measure of the wage rate. There, I used information on average weekly hours and other characteristics from the 1986 Labor Market Activity Survey (LMAS) to impute hours per week by occupation and full-time/part-time status. Average hourly earnings were then constructed from total earnings and from imputed average weekly hours. Here, I prefer to introduce as little noise as possible into the estimation.

dummy for this threshold (T) is thus included in the labor supply equation. The borrowing threshold may not be, however, the only channel through which the debt-service ratio has an effect on the wives' labor supply. Thus, the ratio (X) of housing payments to family income, exclusive of the wife's labor earnings, is also included in the labor supply equation. The rejection of the endogeneity of these variables would be consistent with the view that housing choices are primarily determined by consumption demand for housing rather than investment demand. That is, workers would choose high levels of labor supply because they are burdened by a high debt-service ratio rather than choose high levels of debt-service ratio because they enjoy high levels of labor supply. [17] As the data in Figure 1 shows, at high levels of debt-service ratio, very few married women choose low levels of labor supply.

Since the preceding descriptive analysis also revealed that the 52 weeks in a year constraint was binding for more than 30% of the observations, the two-limit Tobit model will provide the basic workhorse for analysis. [18] A weakness of the standard Tobit model is that it assumes away constraints on minimum amounts of labor supply. Although, this problem may be important in an hours of work equation, it may be less important in a weeks of work equation. As mentioned in Moffitt (1982), the demand side constraints on weeks are not likely to be as severe. Also, the argument that the money costs of labor market participation imply that individuals will not work below some minimum number of hours a week, may hold with less strength when weeks are considered. However, the impact of the time costs of labor market entry or full-time status may still

[17] As shown in Philips and Vanderhoff (1991), the effect of the wife's occupational choice on housing decisions can be captured by her human capital characteristics. It will thus be possible to account for the wife's attachment to the labor market.

[18] Five observations among all homeowners (4 among homeowners with a mortgage) were artificially constrained at 52 weeks, while 747 observations (547) had reached that upper bound.

be important. The fact that almost 10% of wives work part-time full year may indicate that there are constraints not only in labor market entry but also in switching from part-time to full-time status. Yet, the variables that could account for these supply side fixed costs, such as childcare expenses, women's clothing and the number of cars, are likely to be endogenous. Thus in order the simplify the analysis, I prefer to leave these issues for another occasion.

As before, let Y_i denote wife's i choice of number of weeks worked, and let Y_i^* be the latent variable

$$Y_i^* = Z_{1i}'\beta_1 + X_i\beta_2 + T_i\beta_3 + \varepsilon_i \quad i = 1, \ldots, N \qquad (3.12)$$

where the β's are the parameters of the labor supply equation, where Z_{1i}' is a vector of explanatory variables, and the ε_i are independently normally distributed disturbances. The vector Z_{1i}' includes the husband's income, the wife's nonlabor income, the wife's age, the wife's education, the number of children less than seven years old, the number of children between the ages of seven and fifteen, and regional dummies for Quebec, for Ontario, for the Prairies, for Alberta and for British Columbia. The labor supply model is thus given by

$$Y_i = \begin{cases} 52 & \text{if} \quad Y_i^* \geq 52 \\ Y_i^* & \text{if} \quad 0 < Y_i^* < 52 \\ 0 & \text{if} \quad Y_i^* \leq 0 \end{cases} \quad i = 1, \ldots, N. \qquad (3.13)$$

Table VI reports in column (TT) the parameter estimates from the two-limit Tobit maximum likelihood estimation (MLE) of model (3.13) estimated among all homeowners and among homeowners with a mortgage. These estimates show that both the debt-service ratio X and the borrowing threshold T are significant with P-value of 0.015 and 0.000, respectively, among all homeowners and of 0.096 and 0.000, respectively, among homeowners with a mortgage. [19] Furthermore,

[19] By contrast with the descriptive kernel regression of the preceding subsection, here, obser-

the magnitude of the positive effect of the borrowing constraint, which adds approximately 20 weeks of work, makes it comparable to the negative effect of young children, which subtracts approximately 18 weeks of work.

To test the endogeneity of the debt-service ratio (X) and of the borrowing threshold (T), I use a procedure that relies on generalized residuals (as discussed in Pagan and Vella, 1989, Vella, 1993) to perform a specification test in the spirit of Hausman (1978). I also compared the traditional two-stage estimator based on the marginal likelihood (Nelson and Olsen, 1978) of the weeks of work equation with the two-stage estimator based on the conditional likelihood (Smith and Blundell, 1986).

Consider the following reduced forms of X_i and of the latent variable T_i^*

$$X_i = Z_i'\gamma + u_i, \tag{3.14}$$

$$T_i^* = Z_i'\delta + v_i \qquad i = 1, \ldots, N \tag{3.15}$$

where $Z_i' = (Z_{1i}', Z_{2i}')$ and Z_{2i}' is a vector of instrumental variables. These include the ratio of home equity to the value of the home, the square and the cube of this ratio, and dummies for homes built before 1976 and after 1984.[20] Although X_i is continuously observed, observations on T_i^* are dichotomous

$$T_i = \begin{cases} 1 & \text{if} \quad T_i^* > 0 \\ 0 & \text{if} \quad T_i^* \leq 0 \end{cases} \qquad i = 1, \ldots, N. \tag{3.16}$$

The two-stage estimates are obtained by first estimating the γ's via OLS and the δ's with a Probit maximum likelihood procedure, under the assumption that

vations for which $X > 0.5$ are included, but observations where the husband is not working generate outliers which are excluded.

[20] The use of a cubic function of the home equity of a percentage of the value of the home is justified by the fact that the debt-service ratio is likely to be higher for movers and new home-buyers than for other homeowners. An alternative specification which excluded the home equity variables but included the husband's age and education also supported the weak exogeneity of the borrowing threshold, albeit with less strength.

TABLE VI

TOBIT ESTIMATES OF THE WIFE'S NUMBER OF WEEKS WORKED

	All Homeowners			Owners with a Mortgage		
	(TT)	(NO)^b	(SB)^b	(TT)	(NO)^b	(SB)^b
Husband's income (1 000$)	-.3319 ***	-.2711 ***	-.2690 ***	-.3219 ***	-.2205 *	-.2242 *
	(.0595)	(.0751)	(.0855)	(.0693)	(.1336)	(.1490)
Wife's other income (1 000$)	-1.279 ***	-1.205 ***	-1.199 ***	-1.902 ***	-1.711 ***	-1.707 ***
	(.3291)	(.3643)	(.3688)	(.4348)	(.4978)	(.5295)
Wife's age	-1.092 ***	-1.033 ***	-1.038 ***	-.8419 ***	-.8108 ***	-.8138 ***
	(.1215)	(.1350)	(.1356)	(.1405)	(.1567)	(.1573)
Wife's education	2.955 ***	2.849 ***	2.852 ***	2.438 ***	2.303 ***	2.297 ***
	(.3661)	(.4039)	(.4130)	(.3994)	(.4353)	(.4641)
Number of small children (0-6)	-18.75 ***	-18.52 ***	-18.75 ***	-17.34 ***	-17.21 ***	-17.40 ***
	(1.639)	(1.634)	(1.656)	(1.675)	(1.669)	(1.701)
Number of older children (7-15)	-4.612 ***	-4.503 ***	-4.565 ***	-4.577 ***	-4.544 ***	-4.626 ***
	(1.177)	(1.227)	(1.232)	(1.268)	(1.312)	(1.261)
Region of residence^a						
Quebec	3.182	2.083	2.135	2.696	1.101	1.165
	(3.311)	(3.567)	(3.656)	(3.587)	(4.236)	(4.647)
Ontario	11.67 ***	10.78 ***	10.84 ***	9.656 ***	8.306 **	8.351 **
	(3.112)	(3.117)	(3.115)	(3.443)	(3.508)	(3.835)
Prairies	6.888 *	6.292 *	6.303 *	4.359	3.679	3.715
	(3.569)	(3.588)	(3.571)	(3.902)	(3.694)	(4.015)
Alberta	11.16 ***	9.903 **	9.732 ***	10.64 ***	8.814 *	8.567 *
	(3.983)	(3.998)	(4.032)	(4.312)	(4.754)	(5.050)
British Columbia	4.314	3.174	3.024	5.986	4.597	4.361
	(4.117)	(4.235)	(4.224)	(4.534)	(4.906)	(5.117)
Ratio of housing payments to family income, exclusive of wife's labor income(X)	21.20 **	36.82 *	40.02 *	13.55 *	40.54	42.86
	(8.697)	(20.21)	(22.25)	(8.134)	(34.11)	(34.14)
Residual from instrumentation of X^c			-23.16			-31.10
			(38.81)			(47.01)
Dummy T=δ($X \geq$.25)	20.37 ***	24.03 ***	22.44 **	20.24 ***	20.59 **	19.93 **
	(4.324)	(8.217)	(11.12)	(4.034)	(9.267)	(10.93)
Generalised residual from instrumentation of T			-1.511			.0254
			(4.944)			(5.492)
Constant	50.07 ***	44.78 ***	44.91 ***	49.21 ***	42.20 ***	42.52 ***
	(7.375)	(8.311)	(8.325)	(8.022)	(10.69)	(10.59)
σ	43.99	44.36	43.97	40.67	41.14	40.65
	(1.192)	(1.222)	(1.213)	(1.270)	(1.304)	(1.284)
Loglikehood	-6250.40	-6266.97	-6248.80	-4479.96	-4497.11	-4479.41
Number of observations		2417			1682	

Standard errors are in parentheses. The symbol (***), (**) or (*) indicates that the estimate is statistically significant at 1% level, at the 5% level, and at the 10% level, respectively.

^{a} Region omitted is Atlantic Provinces.

...ndard errors calculated by bootstrapping the two-stage procedure with 1000 resamples.

...table VII.

the disturbances v_i are independently normally distributed with variance 1. The results from these estimations are reported in table VII.

The second-stage of the estimation is first performed using a Nelson and Olsen (1978) procedure. Predictions from the first-stage estimations, $\hat{X}_i = Z_i'\hat{\gamma}$ and $\tilde{T}_i = Prob(T_i^* > 0) = \Phi(Z_i'\hat{\delta})$, where $\hat{\gamma}$ is the OLS estimate of γ and $\hat{\delta}$ is the Probit MLE of δ, are used in equation (3.12) in the place of X_i and T_i. Thus the two-limit Tobit MLE is performed with the latent variable [21]

$$Y_i^* = Z_{1i}'\beta_1 + \hat{X}_i\beta_2 + \tilde{T}_i\beta_3 + \epsilon_{1i}. \qquad (3.17)$$

The estimation of limited dependent variable models with endogenous explanatory variables through an instrumental variables procedure has been studied by various authors (Amemiya (1978, 1979), Heckman (1978), Lee (1978, 1979), Nelson and Olsen (1978), and more recently, Newey (1987b) among others). While, under the assumption of normality of the errors, the estimates from two-stage procedures are consistent, they have been shown to be relatively inefficient (Newey, 1987b). In a model where the first-stage estimators are least-squares estimators, Newey (1987b) shows that Amemiya's generalized least squares estimator is asymptotically efficient. Here, because the first stage involves the estimation of a Probit model, his result cannot be used. The one-step or full maximum likelihood estimation would be another efficient alternative. However, since this alternative would be very complicated, it will not be attempted here. Instead, the ineffecient two-stage procedure will be used, but the standard errors will be computed using the bootstrap method (Efron, 1979).

The nonparametric bootstrap is a procedure for estimating standard errors by resampling the residuals and approximating the true distribution of the resid-

[21] The introduction of a non-normal term $\mu = T_i - \tilde{T}_i$ into the errors may lead to small departures from normality.

TABLE VII
ESTIMATES OF THE RATIO (X) OF HOUSING PAYMENTS TO FAMILY INCOME, EXCLUSIVE OF THE WIFE'S AS A FUNCTION OF FAMILY CHARACTERISTICS

	All Homeowners		Owners with a Mortgage	
Dependent Variable:	X	T [a]	X	T [a]
Dwelling built before 1976	-.0212 **	-.0882	-.0256 ***	-.0891
	(.0092)	(.1095)	(.0124)	(.1102)
Dwelling built after 1984	-.0002	.5999 ***	-.0017	.5975 ***
	(.0213)	(.2042)	(.0271)	(.2047)
Home equity as a proportion of value of home owned (HQ)	.1138 **	1.028 *	.1038	.9066
	(.0533)	(.5673)	(.0738)	(.6354)
Home equity squared (HQ^2)	-.1294 ***	-2.124 ***	-.1191 ***	-2.018 ***
	(.0306)	(.3803)	(.0408)	(.4318)
Home equity cubed (HQ^3)	-.1126 ***	-1.489 ***	-.1030 **	-1.370 ***
	(.0309)	(.3919)	(.0453)	(.4613)
Husband's income (1 000$)	-.0023 ***	-.0790 ***	-.0032 ***	-.0784 ***
	(.0002)	(.0058)	(.0003)	(.0058)
Wife's other income (1 000$)	-.0025 **	-.0859 ***	-.0061 ***	-.0862 ***
	(.0012)	(.0272)	(.0023)	(.0277)
Wife's age	.0006	.0085	.0010	.0068
	(.0005)	(.0064)	(.0008)	(.0065)
Wife's education	.0047 ***	.0520 ***	.0054 ***	.0505 ***
	(.0014)	(.0183)	(.0020)	(.0185)
Number of small children (0-6)	-.0024	-.0211	-.0009	-.0233
	(.0063)	(.0709)	(.0084)	(.0710)
Number of older children (7-15)	-.0002	.0953 *	.0008	.1056 *
	(.0047)	(.0569)	(.0065)	.0576)
Region of Residence [b]				
Quebec	.0480 ***	.3133 *	.0579 ***	.2789
	(.0130)	(.1753)	(.0183)	(.1772)
Ontario	.0445 ***	.7350 ***	.0525 ***	.7391 ***
	(.0123)	(.1618)	(.0176)	(.1626)
Prairies	.0246 *	.3552 **	.0259	.3594 **
	(.0142)	(.1814)	(.0202)	(.1824)
Alberta	.0449 ***	.6983 ***	.0581 ***	.7018 ***
	(.0159)	(.1917)	(.0223)	(.1930)
British Columbia	.0458 ***	.9214 ***	.0533 **	.9044 ***
	(.0162)	(.1945)	(.0230)	(.1977)
Constant	.1688 ***	.2449	.1801 ***	.3194
	(.0330)	(.3947)	(.0461)	(.4008)
Loglikehood		-428.01		-419.66
R-square	0.1436		0.0957	
Number of observations	2417		1682	

Note: Standard errors are in parentheses. The symbol (*), (**) or (*) indicates that the estimate is statistically significant
at the 1% level, at the 5% level, and at the 10% level, respectively.
[a] T is a dummy indicating a ratio X greater or equal to 0.25. The estimates are obtained using a probit maximum likelihood procedure.
[b] Region omitted is Atlantic Provinces.

uals by an empirical distribution. Assuming that the parameter estimates are consistent, resampling the residuals generates artificial dependent variables to which the model is refitted. In this artificial environment, the errors in the parameters estimates are observable and the Monte Carlo distribution of these errors can be used to approximate their exact sampling distribution. Emerging theoretical and empirical evidence show that, in many circumstances, the bootstrap method provides estimates as good as conventional asymptotics in large samples (Bickel and Freedman, 1981, Beran, 1982). In particular, in a two-stage least-squares model, Freedman (1984) has shown that the bootstrap will provide correct standard error estimates, even in the presence of heteroskedasticity, under conditions less restrictive than normality of the errors. Further, in small samples, the bootstrap has been shown to actually outperform asymptotics (Freedman and Peters, 1984a, b), as well as commonly used exact tests when the distribution is unknown (Hsu, 1991). [22] The parameters estimates from the Nelson and Olsen procedure, together with bootstrapped standard errors are given in table VI, column (NO). [23] For all variables, with the exception of the variable X, there is very little difference between the point estimates from the non-instrumented and the instrumented estimations.

A formal asymptotically optimal test of weak exogeneity for Tobit models, that is easy to apply, has been proposed by Blundell and Smith (1986). It is a test for the exclusion of the residuals from the instrumental variables regression from the primary equation. This test cannot be applied directly here since one of the variable being instrumented (T_i) is dichotomous. The residuals $v_i(\hat{\delta})$ from

[22] See, Brownstone (1990) for an econometric application of the difficulties linked to the application of the bootstrap.

[23] Very little divergence was found between the incorrect standard errors and the bootstrapped estimates for the N-O procedure. The bootstrap procedure used is the one from Stata version 3.0. One thousand bootstrap replications of the two-stage procedure took between 5 and 7 hours on a 486-50 Mhz IBM PC clone, depending on the number of observations.

(3.15) depend on an unobservable variable T_i^*, so they cannot be used. Following (Gouriéroux *et al.*, 1987, Vella, 1993), the residuals $v_i(\hat{\delta})$ are replaced by their best prediction $\hat{\tilde{v}}_i = \tilde{v}_i(\hat{\delta})$ given that

$$\tilde{v}_i(\delta) = E_\delta[v_i(\delta)\,|\,T_i] = \frac{\phi(Z_i'\delta)}{\Phi(Z_i'\delta)[1-\Phi(Z_i'\delta)]}[T_i - \Phi(Z_i'\delta)] \qquad (3.18)$$

where ϕ and Φ are, respectively, the density and the cumulative density of the standard normal random variable. [24] I then estimate

$$Y_i^* = Z_{1i}'\beta_1 + X_i\beta_2 + \hat{u}_i\lambda + T_i\beta_3 + \hat{\tilde{v}}_i\alpha + \varepsilon_{2i}, \qquad (3.19)$$

using two-limit Tobit MLE. Since the residuals and generalized residuals are generated variables, the correct standard errors are obtained by bootstrapping of the two-stage procedure. The results of this estimation procedure, together with bootstrapped standard errors, are reported in tables VI, column (SB). [25]

The t-statistics of the estimated parameters $\hat{\lambda}$ and $\hat{\alpha}$ of the residuals and generalized residuals from the first stage indicate that the inclusion of these variables in the primary equation is strongly rejected. This result supports not only the weak exogeneity of the borrowing constraint threshold but also that of the ratio of housing payments to family income, exclusive of the wife's labor earnings, albeit less strongly. The sensitivity of the Tobit MLE to departures from the assumption of normality may be a subject of concern. However, given that the point estimates of the parameter λ of the generalized residuals from the instrumentation of the borrowing constraint is close to zero, I expect the weak exogeneity of the borrowing constraint to hold under robust estimation

[24] Monte Carlo evidence in Vella (1993) show that tests of this type have considerably more power than conditional moment tests.

[25] Here, there were significant divergences between the incorrect standard errors and the bootstrapped standard errors for the residual of the variable X, which went from 21.27 to 38.81 among all homeowners, and from 31.41 to 47.01 among homeowners with a mortgage. Other differences were small.

such as the Powell's (1986) symmetrically censored least squares estimator. [26] In Newey (1987a), though the symmetrically censored least squares estimation gave strong evidence of misspecification in the Tobit MLE, it did not reverse the result of the exogeneity test. I would support the exogeneity of the debt-service ratio with less strength.

The exogeneity of nonwife income has been another subject of concern in female labor supply equations. Though it was not rejected in Mroz (1987) using various specifications, it was rejected in Smith and Blundell (1986). Accordingly, I have performed the same analysis as above using a reduced form of the husband's income, where it is replaced by his age, his education and a dummy for the 12 occupations described in the FAMEX. The results of this analysis are given in Tables A-I and A-II in the appendix. The two-limit Tobit MLE produces very comparable estimates for the debt-service ratio and the borrowing threshold. However, with this specification, the instrumentation of the borrowing constraint is not particularly successful, as indicated by the large decrease in the value of the Probit loglikelihood value. Thus, this variable is not longer significant in the two-stage estimations. However, the rejection of the endogeneity of the housing variables is not overturned.

IV. Conclusions

The interaction between mortgage lending rules and female labor supply is important both for policy purposes and for the analysis of labor supply and consumption. In this paper, I use descriptive statistics, kernel regressions and reduced-form analyses to show that mortgage related variables, in particular a variable believed to capture the earnings test of mortgage lending rules, have

[26] Suprisingly, perhaps because of the doubling censoring, the estimated residuals $\hat{\varepsilon}_{2i}$ are not asymmetric. On the basis of skewness, normality of these residuals is not rejected at the 27.35% level. On the basis of kurtosis, it is not rejected at the 4.4% level.

an important impact on the labor supply of married women. The strength of the positive impact of the mortgage qualification constraint on the labor supply of married women exceeds the negative effect of young children. Furthermore, this result cannot be explained by the endogeneity of mortgage choices in an equation of female labor supply.

The implications of this result for the analysis of life-cycle behavior and housing choices are multiples.

1. In the presence of an effective earnings test in the mortgage qualification process, there is a connection between household consumption and income among homeowners with a mortgage. [27] Separability of consumption and leisure in the utility function is not sufficient to insure that such a connection does not exist. In many countries where rates of homeownerships exceed 60%, this connection could be a route via which borrowing constraints could succeed in explaining the "excess sensitivity" of consumption to income, where other liquidity constraints approaches (Hayashi, 1987) have failed.

2. The effect of the earnings test on the labor supply of married women is confirmed to be weakly exogenous. Also, the weak exogeneity of the debt-service ratio cannot be rejected. This is consistent with the view that housing choices are primarily determined by consumption demand for housing rather than investment demand, as shown in (Ioannides and Rosenthal, 1992). Given the lumpiness of housing assets, this can lead to distortions of the households' labor market choices.

3. Since levels of participation in the labor market among of married women are also affected by nonpecuniary variables, such as young children, the existence

[27] This connection may apply for a shorter time period where mortgages of the fixed rate variety.

of an effective borrowing constraint to mortgage choices undermines the user cost approach to housing demand.

Of course, this analysis is limited by the nature of the data. The analysis of the full implications of the earnings test of the mortgage qualification process awaits panel data, which is likely to be available for Canada in less than a decade.

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Appendix

TABLE A-I
TOBIT ESTIMATES OF THE WIFE'S NUMBER OF WEEKS WORKED

	All Homeowners			Owners with a Mortgage		
	(TT)	(NO)[b]	(SB)[b]	(TT)	(NO)[b]	(SB)[b]
Dummy for husband's occupation	Yes	Yes	Yes	Yes	Yes	Yes
Husband's age	.0259	.0876	.0774	-.0284	.0534	.0410
	(.2627)	(.2622)	(.2596)	(.2901)	(.3146)	(.3132)
Husband's education	-1.084 ***	-1.032 **	-1.070 ***	-.9001 **	-.8093	-.8562 *
	(.4331)	(.5108)	(.5304)	(.4667)	(.5858)	(.5621)
Family other income (1 000$)	-.1711	-.0930	-.1111	-.1301	.0066	-.0272
	(.1963)	(.2158)	(.2053)	(.2361)	.3350	(.3107)
Wife's age	-1.201 ***	-1.226 ***	-1.211 ***	-.8816 ***	-.9612 ***	-.9370 **
	(.2656)	(.2647)	(.2639)	(.2951)	(.3343)	(.3256)
Wife's education	2.903 ***	2.861 ***	2.858 ***	2.467 ***	2.391 ***	2.390 ***
	(.4231)	(.4818)	(.4928)	(.4581)	(.5367)	(.5370)
Number of small children (0-6)	-19.15 ***	-19.20 ***	-19.28 ***	-18.19 ***	-18.33 ***	-18.35 **
	(1.659)	(1.858)	(1.905)	(1.701)	(2.043)	(1.968)
Number of older children (7-15)	-5.168 ***	-5.099 ***	-5.185 ***	-5.494 ***	-5.353 ***	-5.461 **
	(1.186)	(1.341)	(1.359)	(1.280)	(1.213)	(1.159)
Region of residence[a]						
Quebec	.7948	-.8584	-.4777	.3055	-2.074	-1.527
	(3.331)	(3.023)	(3.074)	(3.619)	(4.018)	(4.298)
Ontario	9.457 ***	9.183 ***	9.065 ***	7.917 **	7.536 **	7.436 **
	(3.122)	(3.017)	(2.984)	(3.485)	(3.659)	(3.744)
Prairies	5.250	5.118	5.063	2.747	2.954	2.870
	(3.593)	(3.415)	(3.404)	(3.951)	(3.892)	(3.866)
Alberta	9.271 **	9.235 **	8.912 **	8.539 **	8.348 *	8.023 *
	(4.003)	(4.294)	(4.217)	(4.347)	(4.475)	(4.498)
British Columbia	2.950	3.008	2.644	4.013	4.536	4.034
	(4.147)	(4.278)	(4.219)	(4.580)	(4.805)	(4.879)
Ratio of housing payments to family, exclusive of wife's labor income (X)	29.61 ***	70.60 ***	65.70 ***	21.78***	78.16 *	69.88 *
	(8.697)	(26.59)	(25.84)	(8.418)	(51.68)	(47.63)
Residual from instrumentation of X[c]			-40.76			-51.05
			(49.74)			(64.18)
Dummy $T=\delta(X \geq .25)$	22.58 ***	2.563	7.677	22.73 ***	-1.296	4.297
	(4.356)	(15.91)	(15.47)	(4.092)	(22.32)	(20.18)
Generalised residual from instrumentation of T			8.639			10.65
			(8.832)			(12.12)
Constant	47.69 ***	42.14 ***	43.29 ***	38.70 ***	31.45 **	33.09 **
	(11.18)	(11.87)	(11.82)	(12.75)	(18.13)	(18.19)
σ	44.28	45.06	44.26	41.10	42.11	41.08
	(1.200)	(1.185)	(1.157)	(1.285)	(1.355)	(1.370)
Loglikehood	-6260.09	-6292.50	-6259.00	-4491.53	-4524.07	-4490.62
Number of observations		2417			1682	

e: Standard errors are in parentheses. The symbol (***), (**) or (*) indicates that the estimate is statistically significant at 1% level, at the 5% level, and at the 10% level, respectively.
gion omitted is Atlantic Provinces.
ndard errors calculated by bootstrapping the two-stage procedure with 1000 resamples.
e table A-II.

TABLE A-II

ESTIMATES OF THE RATIO (X) OF HOUSING PAYMENTS TO FAMILY INCOME, EXCLUSIVE OF THE WIFE'S AS A FUNCTION OF FAMILY CHARACTERISTICS

Dependent Variable:	All Homeowners X	All Homeowners T^a	Owners with a Mortgage X	Owners with a Mortgage T^a
Dwelling built before 1976	-.0148	.0675	-.0189	.0633
	(.0093)	(.0940)	(.0128)	(.0948)
Dwelling built after 1984	-.0069	.5119 ***	-.0087	.5073 ***
	(.0218)	(.1811)	(.0279)	(.1822)
Home equity as a proportion of value of home owned (HQ)	.1061 **	.7069	.1010	.5547
	(.0545)	(.5114)	(.0758)	(.5732)
Home equity squared (HQ^2)	-.1323 ***	-1.718 ***	-.1286 ***	-1.607 ***
	(.0312)	(.3807)	(.0418)	(.4312)
Home equity cubed (HQ^3)	-.1121 ***	-1.039 ***	-.1060 **	-.8913 **
	(.0316)	(.3591)	(.0466)	(.4268)
Dummy for husband's occupation	Yes	Yes	Yes	Yes
Husband's age	-.0007	.0004	-.0011	-.0013
	(.0011)	(.0110)	(.0015)	(.0112)
Husband's education	-.0025	-.0343 *	-.0033	-.0331 *
	(.0017)	(.0179)	(.0024)	(.0181)
Family other income (1 000$)	-.0020 ***	-.0360 **	-.0031 ***	-.0332 *
	(.0008)	(.0179)	(.0012)	(.0179)
Wife's age	.0007	-.0013	.0012	-.0014
	(.0011)	(.0111)	(.0015)	(.0113)
Wife's education	.0028 *	.0272	.0033	.0249
	(.0017)	(.0179)	(.0024)	(.0181)
Number of small children (0-6)	-.0049	-.1162 **	-.0066	-.1175 *
	(.0065)	(.0617)	(.0086)	(.0620)
Number of older children (7-15)	-.0030	-.0343 *	-.0050	-.0280
	(.0048)	(.0492)	(.0066)	(.0497)
Region of Residence[b]				
Quebec	.0379 ***	.1592	.0454 ***	.1239
	(.0133)	(.1533)	(.0188)	(.1552)
Ontario	.0325 ***	.4729 ***	.0392 **	.4791 ***
	(.0125)	(.1432)	(.0181)	(.1445)
Prairies	.0178	.2995 **	.0173	.2950 *
	(.0145)	(.1597)	(.0207)	(.1611)
Alberta	.0336 **	.4941 ***	.0431 *	.4930 ***
	(.0163)	(.1658)	(.0229)	(.1673)
British Columbia	.0389 **	.6404 ***	.0410 *	.6021 ***
	(.0165)	(.1718)	(.0236)	(.1749)
Constant	.1798 ***	-.8366 *	.1832 ***	-.7227
	(.0478)	(.4987)	(.0701)	(.5079)
Loglikehood		-555.80		-542.86
R-square	0.1173		0.0583	
Number of observations	2417		1682	

Note: Standard errors are in parentheses. The symbol (*), (**) or (*) indicates that the estimate is statistically significant at the 1% level, at the 5% level, and at the 10% level, respectively.

[a] T is a dummy indicating a ratio X greater or equal to 0.25. The estimates are obtained using a probit maximum likelihood procedure.

[b] Region omitted is Atlantic Provinces.

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