Available online at www.sciencedirect.com
SCIENCE DIRECT®
ELSEVIER
Journal of Urban Economics 60 (2006) 50–68
JOURNAL OF Urban Economics
www.elsevier.com/locate/jue

The sensitivity of homeowner leverage to the deductibility of home mortgage interest

Patric H. Hendershott [a,*], Gwilym Pryce [b]

[a] University of Aberdeen Business School, Kings College, Aberdeen AB24 3QY, UK
[b] Department of Urban Studies, University of Glasgow, Glasgow G12 8RS, UK

Received 1 July 2005; revised 3 January 2006

Available online 20 March 2006

Abstract

Mortgage interest tax deductibility is needed to treat debt and equity financing of houses symmetrically. Countries that limit deductibility create a debt tax penalty that presumably leads households to shift from debt toward equity financing. The greater the shift, the less is the tax revenue raised by the limitation and smaller is its negative impact on housing demand. Measuring the financing response to a legislative change is complicated by the fact that lenders restrict mortgage debt to the value of the house (or slightly less) being financed. Taking this restriction into account reduces the estimated financing response by 20 percent (a 32 percent decline in debt vs a 40 percent decline). The estimation is based on 86,000 newly originated UK loans from the late 1990s.

Keywords: Mortgage interest deductibility; Homeowner leverage; Debt tax penalty; Fractional logit regression

In many economies debt financing of housing is penalized relative to equity financing, i.e., interest payments are not fully tax deductible. In the Commonwealth countries—Australia, Canada, New Zealand, and the UK (since 1999)—interest is not deductible at all; in most European countries (the UK in the quarter century prior to 1999) interest is only partially deductible, being limited by a ceiling on the deductible amount, application of a lower tax rate to the deduction, or both. As a result, the Modigliani–Miller [14] debt neutrality theorem does not hold; the user cost of capital for owner-occupied housing (through the weighted average cost of capital) is not

* Corresponding author.
E-mail addresses: phh3939@arvig.net (P.H. Hendershott), g.pryce@socsci.gla.ac.uk (G. Pryce).

0094-1190/$ – see front matter
doi:10.1016/j.jue.2006.01.003

independent of LTV choice.[1] Knowing how LTVs respond to deductibility limits is thus crucial to understanding how housing choices will be affected by changes in such limits, as well as to estimating the revenue gain by the changes.

There are two fundamental problems in explaining LTV behavior. First, the tax penalty on debt usage depends on both intricate tax law provisions and the level of debt usage itself. For example, in the UK during the decade 1983–1992, there was a penalty only for loan amounts above £30,000, and between 1993 and 1999 for some households there was also a penalty for loans below £30,000. In the US, a penalty exists for some low-income households, households with low mortgage debt living in states with low house prices and low taxation, and quite high-income households. Thus estimating the penalty for individual households is complicated. Second, it is almost a universal law that LTVs on newly-originated loans are bounded between (censored at) zero and one (debt is bounded between zero and house value). Moreover, initial LTVs of first-time home buyers are highly skewed toward the highest lender-permitted LTV (Hendershott et al. [12]), while as many as half of older homeowners have zero LTVs (Ling and McGill [13]). The upper bound of unity or the lender-permitted maximum constitutes credit rationing. Because the economic response to a debt penalty is to reduce debt, the more a borrower is rationed, the less his response to a given penalty will be. The combined effect of LTVs suffering from both truncation and a highly skewed distribution creates substantial modeling problems.

This paper addresses these complications and modeling problems. We incorporate the simultaneity of the decision to borrow more or less than the £30,000 threshold by using a pseudo two stage least squares approach, running a probit regression to estimate the probability of being above the ceiling and then using the predicted probability of being above the ceiling to compute the tax penalty on debt faced by households. The fundamental censoring of observed LTVs is addressed by employing censored regression (Amemiya [1]) to model the unrationed response to a tax penalty. These results are then contrasted with responses estimated in credit rationing regime where the LTV is bounded by zero and one. We arrive at the constrained responses by utilizing Papke and Wooldridge's [16] "Fractional Response" estimation.

We employ the UK data set of Hendershott et al. [12]. These data are especially useful because they are a random sample of all house purchasers who made a deliberate leverage decision. Thus we do not need to be concerned with lagged responses of existing mortgagees to changes in the debt penalty and other variables. We illustrate how different estimation methodologies affect estimates of the sensitivity of LTV choice to a debt tax penalty. We then simulate mortgage demand with full deductibility and with zero deductibility and compute the impact of shifting from full to zero deductibility.

The paper is divided into seven sections. We begin with a discussion of the debt tax penalty and some earlier LTV research. Section 2 contains our econometric model. Section 3 discusses estimation alternatives to ordinary least squares. Section 4 presents the data, and Section 5 reports the results. Section 6 uses simulation analysis to compute the percentage decrease in debt usage in response to removal of mortgage interest deductibility, and Section 7 concludes.

1. The debt tax penalty and earlier research on mortgage debt usage

Home mortgage interest deductibility is a means of extending the fundamental tax advantage of owner-occupied housing, the non-taxation of the implicit rents owners pay to themselves and

[1] See Hendershott and Slemrod [9] and Woodward and Weicher [19].

the low (often zero) taxation of capital gains, to households who use debt finance. Deductibility does not make debt cheaper than equity; rather it maintains tax equality between the two costs. Thus to the extent that the interest deduction is limited, there is a tax cost or penalty to using debt and one would expect debt usage to be less.

Mortgage interest has never been *fully* deductible in the US. Low income households or households with low mortgage debt living in states with low house prices and low taxation (state taxes and mortgage interest are the two largest deductible expenses) would select not to itemize expenses because taking the standard deduction instead would lower their taxes (Ling and McGill [13]). For these households there is effectively no deductibility. Further, even if a household did itemize, not all mortgage interest was effectively deductible (the amount of interest that raised total deductible expenses to the standard deduction was "wasted"). The amount of wasted interest (and the number of households that chose not to itemize) grew following the 1986 tax act for two reasons (Hendershott et al. [10]). First, a number of expenses that were previously deductible could no longer be itemized, probably the most important being the interest on consumer credit debt. Second, the standard deduction was sharply increased. Finally, the 1986 act also phased out itemized deductions when household income rose above threshold levels, limiting deductibility for very high income households to as little as 20 percent of their interest paid.

Follain, Ling, and various associates have used the change in the effective deductibility of mortgage interest induced by the 1986 tax act to test the hypothesis that household leverage is sensitive to the tax penalty on debt (Follain and Ling [6], Ling and McGill [13], Follain and Dunsky [5], and Dunsky and Follain [4]). In each case, the leverage of individual households was found to be related significantly to the effective deductibility of mortgage interest. Using the Dunsky and Follain estimates, Follain and Melamed [7] built a simulation model and predicted that removal of the mortgage interest deduction (MID) would lower mortgage debt by 41 percent. We note that this estimate is based on removal relative to 1988 US tax law, not removal relative to full interest deductibility. The latter would be an even larger percentage decline in mortgage debt. This work with US data requires forecasting various unavailable household expenses and determining whether households would itemize or take the standard deduction. Moreover, the empirical analysis is not of households at their decision point (when the loan is originated), but wherever they happen to be in the debt cycle, including whether they have a below-market, fixed-rate mortgage (which would obviously dampen borrower incentives to reduce debt).

In general, the weighted average cost of capital for owner-occupied housing is just an average of the debt (*CD*) and equity (*CE*) costs where the weights are the loan to value ratio, *LTV*, and $1 - LTV$:

$$WACC = (LTV)CD + (1 - LTV)CE. \tag{1}$$

If both costs equal the after-tax interest rate, $(1 - T)r$ (we abstract from risk premia), then $WACC = (1 - T)r$. However, if a tax penalty at rate p is imposed per unit of interest paid, the cost per unit of debt is $(1 - T + p)r$ and

$$WACC = (1 - T)r + LTVrp. \tag{2}$$

If the penalty is the nondeductibility of interest, $p = T$ and the *WACC* is increased by the product $LTVTr$.

How much removal of the MID would raise the WACC (and tax revenue) and act to reduce housing demand depends on how much households change their LTV in response to the loss of deductibility. The more households reduce their LTVs, the less the WACC is increased and thus the less will be the reduction in homeownership and housing demand. Also, the less revenue the government would gain by removing the MID.

UK borrowers have been subject to substantially greater variation in limitations on interest deductibility than US borrowers. During the last quarter century, the mortgage interest deduction in the UK has been limited in two ways. First, in 1974 the deduction was restricted to that on a £25,000 mortgage (and the deductibility of interest on other household debt was eliminated). In 1983, the limit was raised to £30,000 (the median UK house price level had nearly tripled to £29,400). Subsequently the limit was never again raised in spite of rising house prices (the median tripled again to £87,300 in 1999). Second, the maximum tax rate at which interest could be deducted was cut from the 40 percent maximum income tax rate to 25 percent in 1992, to 20 percent in 1994, and to 10 percent in 1995 (finally to zero in 1999). Given that there were effectively only two household income tax brackets during this period, 25 and 40 percent, after 1993 no household paying taxes could deduct mortgage interest at their full marginal income tax rate.

Of the ceiling and tax rate maximums, the former has been far more important for new borrowers who have reasonably high initial loan-to-value ratios (the average of our sample is 0.78). With a median house price in 1995–1998 of £53,000 outside the London/Southeast region and £76,000 within this region, over four-fifths of new mortgage originations were above the £30,000 mortgage limit and thus interest was not deductible at all on the margin.

Figure 1 illustrates how the tax penalty per unit debt varies with loan size. The product of the tax rate and a given interest rate are on the vertical axis and loan amount is on the horizontal. The solid line is a household's marginal tax rate times the given interest rate. Holding house value constant, the larger is the mortgage loan, the more interest is deductible (unless the loan is above the ceiling, L_c), but the larger are the household's taxable investments and thus the higher is its taxable interest income. As long as the ceiling is not binding, taxable income is at least roughly

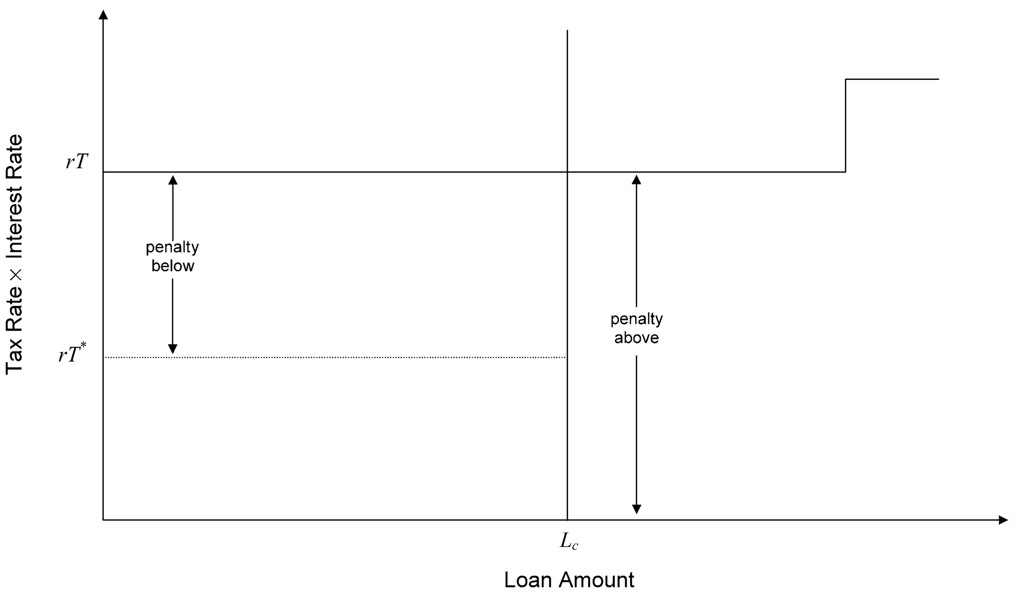

Fig. 1. The tax penalty per unit of debt.

independent of loan size.[2] The dashed rT^* line in the figure is the product of the given interest rate and the lower maximum tax rate, introduced in 1993, at which mortgage interest on the loan amount below the ceiling could be deducted. The tax penalty for loans above L_c is the product of the household's tax rate and the interest rate (Tr). The penalty below the ceiling is the product of the interest rate and the maximum of $T - T^*$ or zero (the latter for households with $T < T^*$). The difference $T - T^*$ is independent of loan size.[3]

Let ψ_i be the probability that the ith borrower's loan exceeds the tax deductibility ceiling of £30,000. The tax penalty per unit of interest paid is then

$$p_i = p_i^A + p_i^B = \psi_i T_i + (1 - \psi_i) \max\left[T_i - T^*, 0\right], \tag{3}$$

where the A and B refer to penalties applying to loans above and below the loan ceiling. Hendershott et al. [12] estimated ψ for individual households and then made the log of *LTV* a function of the two tax penalty variables, p^A and p^B. The log of *LTV* was significantly negatively related to both.

There are two problems with that treatment of the tax penalty. First is the obvious negative correlation of the two variables owing to the complementary of the ψ and $1 - \psi$ terms. Second is the use of predicted loan size (above or below £30,000) to explain LTV. Smaller loans are likely to carry lower LTVs and thus the estimated negative impact of the penalty on small loans will be biased upward relative to that on large loans. In this paper, we compute individual household p_is from Eq. (3) and use a single tax penalty variable, rp_i, in the estimation.

2. Econometric model

A household's demand for mortgage debt is driven by its consumption demand for housing relative to its wealth and the tax penalty. At one extreme are households with sufficient wealth relative to housing consumption demand that they demand zero mortgage debt. At the other extreme are households with a sufficiently high consumption demand relative to wealth that they select the maximum debt allowed by lenders (Hendershott et al. [11]). In the more general case, households are trading off the debt penalty against portfolio diversification considerations (Brueckner [3]). More debt is relatively expensive, but it allows the holding of additional nonhousing assets.

Given the constraints of zero and unity (or slightly less) on the LTV and its importance to the WACC, the LTV is a natural focal point for investigation. However, because the demand for debt is not necessarily homogeneous of degree one in housing demand, determinants of this demand, such as real income and real house prices, are relevant to LTV choice.[4] Other determinants are variables that place households on the need to diversify spectrum. In the absence of a wealth measure, these include borrower age and whether or not they previously owned their home.

[2] When the loan exceeds the ceiling, the larger is the loan (and thus taxable investments), the higher is the household's taxable income. Thus at some point the household could be pushed into a higher tax bracket.

[3] These tax penalties, Tr and $(T - T^*)r$, also apply to some US borrowers, but the application is much different. Whereas the Tr penalty applies to large loans in the UK, in the US it applies only if the loan is sufficiently small that households minimize their tax liability by taking the standard deduction instead of itemizing. And whereas the $(T - T^*)r$ penalty applies to small loans in the UK, in the US it exists only for households with income sufficiently high that interest expense is not fully deductible.

[4] Most of the recent literature explains the quantity of mortgage debt and housing demand in a simultaneous equation framework (Follain and Dunsky [5], and Ling and McGill [13]). Our formulation avoids the necessity of estimating the housing demand relationship.

The diversification benefits of a higher LTV decrease with the level of the LTV; the lower the LTV, the more diversified the household would be. Thus the response of households to the debt penalty, the cost of greater diversification, should be greater the lower the LTV would be in the absence of the penalty. For example, the response should be greater for previous owners and for those in areas with lower real house prices. As a result, we partition our sample along these lines to see how the response to the penalty varies.

Following the above discussion, LTV_i depends on basic or wage income (Y_i^B), other income (Y_i^O), the borrower's age (AGE_i), whether the borrower was a previous owner (PO_i), the debt tax penalty (rp_i) and a vector of location and year dummy variables ($\mathbf{DUM}_i$). Thus the structural equations of the econometric model are given by Eq. (3) above and Eq. (4),

$$LTV_i = LTV_i\big(PO_i, Y_i^B, Y_i^O, AGE_i, rp_i, \mathbf{DUM}_i\big), \tag{4}$$

where

$$\psi_i = \psi_i\big(PO_i, Y_i^B, Y_i^O, AGE_i, LTV_i, \mathbf{DUM}_i\big), \tag{5}$$
$$T_i = T_i, \big(Y_i^B, Y_i^O\big),$$
$$T^* = \text{the maximum marginal tax rate at which interest can be deducted (equal to 0.1 during the period under consideration).}$$

Substituting into Eq. (3), the relationship between p_i and its underlying variables is:

$$p_i = f(\psi_i) = f\big(PO_i, Y_i^B, Y_i^O, AGE_i, r, LTV_i, T^*, \mathbf{DUM}_i\big). \tag{3$'$}$$

That is, p_i depends on ψ_i which in turn depends on LTV_i. We control for the dependency of ψ_i on LTV_i by using $\psi_i^{\#}$, the predicted values from a probit regression of ψ_i on all exogenous variables in the system (i.e., all variables excluding LTV—see Greene [8, p. 398]). We then compute the $p_i^{\#}$ from

$$p_i^{\#} = \psi_i^{\#} T + \big(1 - \psi_i^{\#}\big) \max[T_i - T^*, 0] \tag{3$''$}$$

and substitute these for the p_i in Eq. (4). Finally we estimate the LTV response.[5]

3. Estimation alternatives

Our econometric estimation is complicated by the possible credit rationing that arises from lenders placing an upper bound on LTV at (or near) one. In perfectly informed credit markets we would not expect either equilibrium credit rationing or collateral constraints of the kind described respectively by Stiglitz and Weiss [17] and Bester [2], and OLS regression would provide unbiased estimates. But this is not the world we live in.

A likely response for households where the benefits of owning with unrestricted lending over renting are large is to take the maximum LTV on offer, even if this is below their optimum.[6] If

[5] In a world where mortgage interest is not tax deductible, the marginal tax rate paid by a household can depend on the LTV chosen. To illustrate, for a given house purchase, the larger the loan chosen, the more wealth the household has to invest in other assets and thus the higher is the taxable income of the household. To eliminate this source of endogeneity, we add an estimate of the interest a household could earn on their equity investment in the house to the income of the household before computing their marginal tax rate. The estimate is computed as the product of the mortgage rate and the difference between house purchase price and mortgage loan amount.

[6] Some households might choose to delay homeownership for a period.

this were the case and rationing were common, we would expect to see many observations with LTVs exactly equal to the maximum. If the latent variable y^* is observed if $y^* < y'$ and is not observed if $y^* \geqslant y'$, then the observed dependent variable is defined as:

$$y = \begin{cases} y_i^* = \beta x_i + u_i & \text{if } y_i^* < y', \\ 1 & \text{if } y_i^* \geqslant y', \end{cases}$$

where $u_i \sim \mathrm{IN}(0, \sigma^2)$. This is the tobit model, first suggested by Tobin [18].

The basic tobit model has limited usefulness when applied to our data because less than one percent of observations (850 cases out of 86,620 records) have LTVs exactly equal to one (when we applied tobit to our data, coefficients and predicted values unsurprisingly were virtually identical to OLS estimates). Note that the low incidence of observations with LTVs exactly equal to one is not necessarily indicative of the absence of censoring. While some lenders may set the global maximum LTV at unity, others will set a maximum below this level, possibly based on their perception of the creditworthiness of the borrower concerned. Indeed, we find two lower cut-points, 0.95 and 0.90, with 15,254 and 3848 cases having LTVs exactly equal to these thresholds, respectively. Together these cut-points represent 22 percent of our total sample. If a significant proportion of these observations are the result of credit rationing, rather than because borrowers choose to be at these points owing to the piecewise nature of mortgage pricing, then we need to account for the censoring of the dependent variable.[7] We do this by applying Amemiya's [1] Censored Regression model (CR) which is a generalized version of tobit regression that allows for variable cutoffs (in our case, at LTVs of 1.00, 0.95, and 0.90).

While Censored Regression will help us ascertain what the parameters of our model would be in the absence of rationing (and hence predict what the unconstrained LTV would be), like OLS the predicted CR values can exceed one. A popular way of modeling variables bounded between zero and one is to apply the log-odds transformation to the dependent variable ($\log[y/(1-y)]$) which allows OLS to be applied to the estimation of $\mathbf{x\beta}$. This approach has two major drawbacks for our estimation of LTVs:

> "First, it cannot be used directly if y takes on the boundary values, zero and one. While we can always use adjustments for the boundary values, such adjustments are necessarily arbitrary. Second, even if y is strictly inside the unit interval, $\boldsymbol{\beta}$ is difficult to interpret: without further assumptions, it is not possible to recover an estimate of $\mathrm{E}(y|\mathbf{x})$, and with further assumptions, it is still nontrivial to estimate $\mathrm{E}(y|\mathbf{x})$." (Wooldridge [20, p. 662])

The solution suggested by Papke and Wooldridge [16] and Wooldridge [20] is to model $\mathrm{E}(y|\mathbf{x})$ as a logistic function:

$$\mathrm{E}(y|\mathbf{x}) = \exp(\mathbf{x\beta})/\big[1 - \exp(\mathbf{x\beta})\big].$$

This method guarantees that "predicted values for y are in (0,1) and that the effect of any x_i on $\mathrm{E}(y|\mathbf{x})$ diminishes as $\mathbf{x} \to \infty$" (Wooldridge [20, p. 662]). Beyond the bias issue, the high LTVs of many borrowers in our sample even in the face of a tax penalty implies that removing the tax penalty will likely predict LTVs over 100 percent for many borrowers. Thus more sophisticated estimation techniques are needed if we want to estimate the constrained response.

[7] There is no evidence of a straightforward positive relationship between the interest rates charged on mortgage contracts and LTVs. Indeed, the average rate for 95 percent LTVs is a half basis point less than that for 90 percent LTVs.

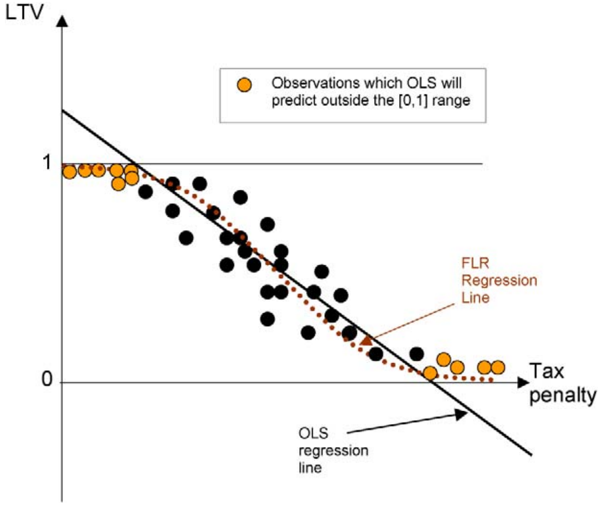

Fig. 2. Constrained response: predicting with the [0,1] interval.

Moreover, simulated increases in LTV owing to removal of the tax penalty will also be constrained not to exceed unity. The effect is illustrated in Fig. 2 where it can be seen that OLS can simulate outside the [0,1] range at the extremes, whereas the recumbent "S" shape of the FLR fitted line ensures that the LTV constraints are binding for predicted values. Because the LTV regression is the second part of a two-stage system and the right-hand-side endogenous variable (the debt tax penalty) enters non-linearly, the usual 2SLS standard errors are unlikely to apply to our OLS estimates. Thus we report the bootstrapped standard errors (based on 1000 repetitions) for all regression results, rather than the 2SLS standard errors or those suggested by Papke and Wooldridge in the single equation context for FLR.[8]

4. Data

The empirical analysis is based on over eighty-six thousand observations extracted from the Council for Mortgage Lenders' annual survey of 5% of all mortgage originations over the period 1995 to 1998 (see Hendershott et al. [12] for details). Because variable descriptives and regression coefficients tended to be noticeably different for London and the Southeast compared with the remaining Other regions, we present two sets of results in each table. Variable definitions and descriptives are presented in Table 1 (i) for London/SE and in Table 1 (ii) for Other. If a variable is continuous, we present the mean, median, and coefficient of variation (the standard deviation as a proportion of the mean); if a variable is dichotomous, we present the proportion. Notice that the mean house price is almost 50 percent higher in the London/SE region than elsewhere, and income is 30 percent higher. As a result of the former, the proportion of loans above £30000 is much greater, 90 percent in London/SE versus 77 percent in Other. Half of all loans have LTVs of 89 percent or higher.

[8] See Mooney and Duval [15] for an introduction to bootstrapping.

Table 1
Descriptive statistics

Variable	Description	(i) London/SE $N = 33{,}541$			(ii) Other regions $N = 53{,}079$		
		Mean or proportion	Median	CV	Mean or proportion	Median	CV
Age	Age of main borrower	34.65	32.00	0.08	34.50	32.00	0.08
LTV	Loan to value ratio	0.77	0.87	0.08	0.79	0.90	0.07
Income	Basic income at 1990 prices (£000s)	15.94	13.01	0.49	12.03	10.13	0.42
Other income	Other income at 1990 prices (£000s)	4.10	0.00	2.47	3.18	0.00	2.20
Mortgage	Amount borrowed (£000s)	67.47	59.00	0.34	47.37	42.00	0.29
Price	Purchase price (£000s)	94.28	76.00	0.45	63.87	53.75	0.37
Dummies:							
ceil	= 1 if mortgage amount is > £30K	0.90	–	–	0.77	–	–
prev_oo	= 1 if a previous owner	0.54	–	–	0.51	–	–
age_lt25	= 1 if age < 25	0.10	–	–	0.13	–	–
age25_34	= 1 if aged 25 to 34	0.48	–	–	0.46	–	–
age35_44	= 1 if aged 35 to 44	0.26	–	–	0.25	–	–
age45_54	= 1 if aged 45 to 54	0.12	–	–	0.12	–	–
age_gt54	= 1 if age > 54	0.04	–	–	0.04	–	–
incoth_d	= 1 if income other than basic is > 0	0.46	–	–	0.48	–	–
1996	= 1 if mortgage taken out in 1996	0.25	–	–	0.24	–	–
1997	= 1 if mortgage taken out in 1997	0.31	–	–	0.29	–	–
1998	= 1 if mortgage taken out in 1998	0.23	–	–	0.24	–	–
LondSE	= 1 if located in London/S.East	1.00	–	–	0.00	–	–
Y.Hum	= 1 if located in Yorks/Humberside	0.00	–	–	0.14	–	–
E.Mids	= 1 if located in East Midlands	0.00	–	–	0.13	–	–
E.Ang.	= 1 if located in East Anglia	0.00	–	–	0.07	–	–
Lon.	= 1 if located in London	0.34	–	–	0.00	–	–
S.East	= 1 if located in South East	0.66	–	–	0.00	–	–
S.West	= 1 if located in South West	0.00	–	–	0.17	–	–
W.Mids	= 1 if located in West Midlands	0.00	–	–	0.14	–	–
N.West	= 1 if located in North West	0.00	–	–	0.16	–	–
Scot.	= 1 if located in Scotland	0.00	–	–	0.11	–	–

Table 2 divides the two area samples by age, buyer status (first time owners, FTOs, and previous owners, POs), and the tax rate paid, either 0.23 to 0.25 or 0.4. Of course, FTOs are younger than POs; nearly four-fifths of FTOs are under age 35, while only two-fifths of POs are. Conversely, less than a tenth of FTOs are over age 44, versus a quarter of POs. Holding age constant, POs have higher incomes than FTOs and are thus more likely to be paying the higher 40 percent tax rate. This is especially true because the POs, being older and having higher incomes, have greater imputed income on equity invested in the house. These statements are equally valid for those in the London/SE as those living elsewhere.

The mean LTVs for each age/tax-rate cell are also listed. As expected, LTVs decline with age and are lower for POs. The age impact is far less for FTOs than POs; LTVs decline by 0.17 (London/SE, low tax rate) and 0.20 (Other, low tax rate) between under age 25 to over 54 for FTOs versus declines of 0.35 (London/SE) and 0.39 (Other) for POs. Finally, as expected, LTVs are lower for borrowers in the higher (0.4) tax bracket, although this is not true for the oldest PO households.

Table 2
Cell counts and average LTVs

Age:		(i) London/SE: First time owners tax rate: 0.23–0.25	0.4	Total	Previous owners tax rate: 0.23–0.25	0.4	Total	(ii) Other regions: First time owners tax rate: 0.23–0.25	0.4	Total	Previous owners tax rate: 0.23–0.25	0.4	Total
< 25	$\overline{LTV}$	**0.91**	**0.87**		**0.76**	**0.64**		**0.92**	**0.89**		**0.82**	**0.77**	
	n	2802	152	2954	434	66	500	6156	70	6226	654	43	697
	n as %	18.1	1.0		2.4	0.4		23.8	0.3		2.4	0.2	
25–34	$\overline{LTV}$	**0.90**	**0.88**		**0.79**	**0.73**		**0.90**	**0.87**		**0.81**	**0.76**	
	n	7114	1870	8984	4119	3081	7200	13,043	881	13,924	8178	2086	10,264
	n as %	45.9	12.1		22.8	17.1		50.5	3.4		30.0	7.7	
35–44	$\overline{LTV}$	**0.86**	**0.85**		**0.69**	**0.66**		**0.86**	**0.84**		**0.71**	**0.68**	
	n	1695	838	2533	2610	3442	6052	3233	649	3882	6062	3218	9280
	n as %	10.9	5.4		14.5	19.1		12.5	2.5		22.2	11.8	
45–54	$\overline{LTV}$	**0.82**	**0.79**		**0.56**	**0.55**		**0.81**	**0.79**		**0.58**	**0.57**	
	n	586	253	839	1552	1571	3123	1172	258	1430	3321	1787	5108
	n as %	3.8	1.6		8.6	8.7		4.5	1.0		12.2	6.6	
> 54	$\overline{LTV}$	**0.74**	**0.73**		**0.41**	**0.47**		**0.72**	**0.67**		**0.44**	**0.50**	
	n	140	36	176	801	379	1180	322	37	359	1491	418	1909
	n as %	0.9	0.2		4.4	2.1		1.2	0.1		5.5	1.5	
Totals:		79.7%	20.3%	15,486	52.7%	47.3%	18,055	92.7	7.3%	25,821	72.3%	27.7%	27,258
							33,541						53,079

Note. $\overline{LTV}$ represents the mean loan to value ratio in a cell.

5. Equation estimates

The probit equations for loans above £30,000 were estimated separately for the London/SE and the Other regions and are reported in Appendix Table. The coefficients from these regressions were used to predict the probability of a household's loan exceeding the £30,000 ceiling, and this probability was used to compute the tax penalty per unit of interest paid, $p_i^{\#}$, from Eq. ($3''$). The tax penalty variable employed in the LTV equations is the product of these and r.

Table 3 presents the LTV regressions for previous owners and first-time owners using the different regression techniques. The tax penalty variable is highly statistically significant (t-ratios, in parentheses, range from 10 to 35), irrespective of the regression technique (OLS, CR, and FLR) employed, as are basic income, other income, and age. As for the OLS and CR tax-penalty coefficients, one should anticipate that the OLS coefficient would be biased towards zero due to the omission of totally constrained borrowers from the CML data (i.e., households who were unable to obtain a mortgage at the minimum LTV needed to purchase a dwelling that would yield greater utility than their current dwelling). This is confirmed: the OLS coefficients are smaller in absolute terms than the CR coefficients in both the previous owner and first time buyer regressions for both the London/SE and Other regions (and are significantly smaller in three of four cases). Note also that the tax-penalty responses under the OLS and CR estimations for first-time buyers are roughly half those for previous owners. This is expected given the younger age, and thus lower accumulated wealth, of first-time buyers.

Because the FLR coefficients are not the first partial derivatives (due to the non-linear structure of the logit functional form), they cannot be directly compared with the OLS and CR coefficients.

Table 3
LTV regressions

	FTO			PO		
	OLS	CR	FLR	OLS	CR	FLR
(i) London/SE						
$rp^{\#}$	−5.596	−7.163	−52.186	−11.400	−12.485	−61.891
	(−14.9)	(−15.3)	(−16.1)	(−11.5)	(−11.6)	(−12.0)
inc.	0.012	0.014	0.113	0.018	0.019	0.108
	(8.7)	(9.6)	(8.2)	(5.2)	(5.2)	(5.6)
inc.^{2}	−2.0E–04	−2.4E–04	−2.0E–03	−2.3E–04	−2.4E–04	−1.6E–03
	(−4.3)	(−4.8)	(−4.2)	(−2.9)	(−2.9)	(−3.3)
inc.^{3}	8.7E–07	1.0E–06	9.8E–06	6.7E–07	7.1E–07	5.9E–06
	(2.3)	(2.5)	(2.4)	(1.5)	(1.5)	(2.0)
O.inc.	0.005	0.007	0.052	0.010	0.011	0.053
	(10.2)	(8.5)	(8.7)	(7.6)	(7.8)	(7.2)
O.inc.^{2}	−2.9E–04	−3.6E–04	−2.7E–03	−3.5E–04	−3.8E–04	−1.8E–03
	(−6.2)	(−5.0)	(−4.7)	(−3.4)	(−3.5)	(−3.2)
O.inc.^{3}	2.6E–06	3.4E–06	2.6E–05	2.2E–06	2.4E–06	1.1E–05
	(4.4)	(3.0)	(2.7)	(1.5)	(1.6)	(1.5)
age	−0.001	−0.002	−0.036	−0.009	−0.012	−0.066
	(−0.6)	(−1.9)	(−4.7)	(−7.7)	(−9.9)	(−10.0)
age^2	−4.5E–05	−3.3E–05	4.1E–05	−8.0E–06	1.5E–05	2.1E–04
	(−3.0)	(−1.9)	(0.4)	(−0.5)	(1.1)	(2.9)
1996	−0.017	−0.021	−0.146	−0.045	−0.046	−0.249
	(−6.3)	(−5.4)	(−5.5)	(−6.7)	(−6.8)	(−7.8)
1997	−0.024	−0.029	−0.224	−0.040	−0.043	−0.212
	(−9.0)	(−8.3)	(−8.5)	(−7.4)	(−8.0)	(−8.3)
1998	−0.056	−0.060	−0.510	−0.041	−0.043	−0.213
	(−18.5)	(−14.8)	(−18.7)	(−7.9)	(−7.9)	(−8.6)
S.East	0.009	0.013	0.091	0.029	0.033	0.141
	(4.9)	(4.7)	(5.0)	(8.2)	(8.8)	(8.4)
Const.	0.953	1.033	3.135	1.060	1.152	2.970
	(53.3)	(45.0)	(22.3)	(44.1)	(44.5)	(19.2)
N	15,486	15,486	15,486	18,055	18,055	18,055
Adj. R^2	0.122	–	–	0.236	–	–
LL	11,624	1367	−4108	2531	−1188	−7953
AIC	−23,200	−2705	8244	−5033	2406	15,933
(ii) Other regions						
$rp^{\#}$	−5.911	−7.508	−53.622	−13.471	−14.921	−68.923
	(−12.36)	(−12.29)	(−13.99)	(−34.92)	(−34.57)	(−36.88)
inc.	0.009	0.011	0.080	0.031	0.033	0.161
	(5.47)	(4.92)	(4.42)	(23.52)	(22.79)	(36.08)
inc.^{2}	−1.1E–04	−1.3E–04	−9.2E–04	−5.5E–04	−6.0E–04	−3.0E–03
	(−1.66)	(−1.41)	(−1.26)	(−13.31)	(−13.00)	(−23.13)
inc.^{3}	2.9E–07	3.4E–07	2.3E–06	2.7E–06	2.9E–06	1.5E–05
	(0.58)	(0.44)	(0.42)	(8.29)	(8.24)	(15.38)
O.inc.	0.008	0.010	0.090	0.012	0.012	0.067
	(10.95)	(8.57)	(12.08)	(8.22)	(8.37)	(9.78)
O.inc.^{2}	−5.8E–04	−6.9E–04	−6.3E–03	−4.2E–04	−4.5E–04	−2.8E–03
	(−5.86)	(−4.60)	(−6.45)	(−3.30)	(−3.31)	(−4.39)
O.inc.^{3}	6.7E–06	8.1E–06	8.3E–05	3.5E–06	3.7E–06	2.7E–05
	(3.13)	(2.34)	(3.41)	(1.79)	(1.76)	(2.63)

(*continued on next page*)

Table 3 (*continued*)

	FTO			PO		
	OLS	CR	FLR	OLS	CR	FLR
age	−0.002	−0.004	−0.049	−0.017	−0.022	−0.112
	(−2.10)	(−4.03)	(−9.08)	(−19.02)	(−21.73)	(−24.81)
age^2	−3.7E–05	−2.3E–05	1.4E–04	6.8E–05	1.1E–04	6.7E–04
	(−3.08)	(−1.68)	(1.96)	(6.32)	(9.21)	(12.92)
1996	−0.016	−0.016	−0.140	−0.044	−0.046	−0.226
	(−6.95)	(−4.93)	(−5.67)	(−11.23)	(−10.63)	(−11.02)
1997	−0.017	−0.016	−0.154	−0.025	−0.024	−0.119
	(−7.92)	(−5.23)	(−6.91)	(−6.91)	(−6.00)	(−6.45)
1998	−0.040	−0.039	−0.383	−0.012	−0.007	−0.054
	(−18.25)	(−12.41)	(−17.18)	(−3.28)	(−1.71)	(−2.84)
Y.Hum	0.006	0.006	0.064	0.003	0.003	0.017
	(2.09)	(1.29)	(2.04)	(0.59)	(0.45)	(0.58)
E.Mids	8.5E–04	−4.3E–03	7.7E–03	4.7E–04	1.3E–04	5.0E–03
	(0.27)	(−0.98)	(0.24)	(0.08)	(0.02)	(0.17)
N.West	0.010	0.005	0.109	0.002	0.001	0.011
	(3.46)	(1.21)	(3.59)	(0.32)	(0.09)	(0.37)
Scot.	−0.002	−0.007	−0.027	0.007	0.004	0.039
	(−0.64)	(−1.44)	(−0.76)	(1.32)	(0.63)	(1.36)
W.Mid	−0.002	−0.011	−0.020	−0.011	−0.013	−0.052
	(−0.54)	(−2.45)	(−0.62)	(−1.99)	(−1.98)	(−1.80)
S.West	−0.005	−0.013	−0.051	−0.006	−0.008	−0.027
	(−1.57)	(−2.79)	(−1.54)	(−1.22)	(−1.36)	(−0.97)
E.Ang.	−0.006	−0.012	−0.065	0.001	0.003	0.013
	(−1.64)	(−2.24)	(−1.76)	(0.23)	(0.36)	(0.38)
Const.	1.007	1.116	3.718	1.205	1.333	3.896
	(66.99)	(54.02)	(28.52)	(61.80)	(62.62)	(38.64)
N	25,821	25,821	25,821	27,258	27,258	27,258
Adj. R^2	0.123	–	–	0.277	–	–
LL	19,697	1329	−6641	4681	−1876	−11,600
AIC	−39,400	−2617	13,323	−9322	3795	23,161

Notes. All figures in parentheses are t-ratios derived from bootstrapped standard errors using 1000 repetitions. "LL" stands for log-likelihood; $rp^{\#}$ is the product of market interest rate r and the estimated tax penalty $p^{\#}$.

We note, though, that here, too, the responses of previous owners are significantly greater than those of first-time owners.

Figure 3 illustrates the general fit of the equations by comparing the kernel density function of the predicted values of our three estimation techniques with the kernel density plot of the actual values of LTV.[9] Note that none of the estimation techniques explains the extreme concentration of LTVs at or above 0.95 (almost a third of our sample), although the CR method seems to do a marginally better job. On the other hand, the CR method predicts a number of LTVs greater than unity. The OLS and FLR methods have quite similar distributions of predicted values.

[9] "Kernel density estimation" is the term used to describe a series of algorithms for approximating the density function of an observed distribution. We use the standard Epanechnikov kernel with a half-width (the width of the density window around each point) set at 0.01 for Figs. 3 and 4.

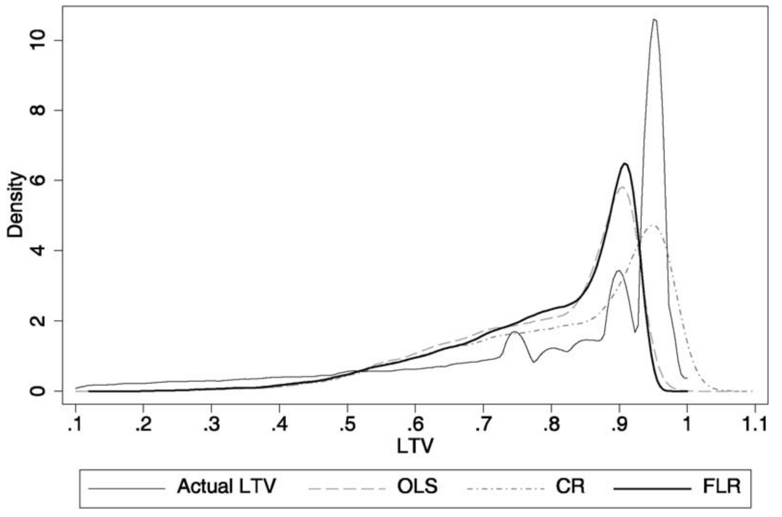

Fig. 3. Kernel density plots: actual LTV vs. predicted.

6. Simulated impact of removing the mortgage interest deduction (MID)

During our estimation period all borrowers effectively had partial interest deductibility. To compute the impact on debt usage of shifting from full deductibility to no deductibility (or vice versa) then requires two simulations: one to determine what LTVs would be with full deductibility (no tax penalty) and another to determine LTVs with zero deductibility (tax penalty equal to Tr). The percentage reduction in LTV owing to the loss of interest deductibility is then computed as the difference between the latter simulated value and the former, times 100, divided by the former.

Table 4 reports mean percentage declines in LTV resulting from removal of the MID implied by each of the three estimation methods for each of the cells in Table 2. The percentage decline is greatest (across all estimation methods) for older borrowers (particularly previous owners) because their LTV is lower at the outset (full deductibility) so the same absolute fall in LTV will have a larger percentage effect. The percentage declines for previous owners are roughly double those for first time owners (less than double for young households according to the FLR estimation and somewhat more for older households according to OLS estimation). This follows directly from the larger tax penalty coefficients in Table 3. Simulated percentage decreases are also greater for households in the 40 percent tax bracket than those in the 23–25 percent brackets (not that much greater for older households). This is again as expected because the former are having a larger tax penalty removed. Finally, simulated decreases are quite similar for the London/SE and Other regions.

Regarding the estimation method, the CR method consistently has estimates 20 percent higher than OLS for FTOs and 5 to 10 percent higher for POs, which is consistent with the argument that the OLS estimates are biased downward. FLR, on the other hand, gives smaller percentages than OLS or CR for those under age 44 and larger increases for older borrowers (larger than CR only for households over age 54). The reason for the smaller FLR simulated impacts for younger borrowers is clear from Fig. 4, which gives the simulated LTV distributions for the three estimation methods with full interest deductibility. The OLS and CR simulations have many borrowers

Table 4
Percentage decline in LTV due to a shift from full to zero interest deductibility

		London/SE				Other regions			
		FTO		PO		FTO		PO	
Age:		0.23–0.25	0.4	0.23–0.25	0.4	0.23–0.25	0.4	0.23–0.25	0.4
< 25	OLS	10.2	16.9	20.1	31.6	10.6	17.2	21.9	33.3
	CR	12.2	19.9	20.9	32.9	12.5	20.1	22.8	34.5
	FLR	7.3	11.4	12.8	20.2	6.7	9.5	10.7	14.7
25–34	OLS	10.3	17.1	21.0	32.8	10.8	17.7	23.0	35.2
	CR	12.3	20.2	21.9	34.3	12.7	20.8	24.0	36.7
	FLR	8.0	12.9	15.3	23.4	7.9	12.5	13.9	20.2
35–44	OLS	10.7	17.5	23.4	35.4	11.2	18.3	25.8	38.3
	CR	12.9	20.8	24.7	37.3	13.4	21.7	27.3	40.3
	FLR	10.8	15.7	22.1	30.4	11.1	15.8	21.5	29.1
45–54	OLS	11.5	18.5	27.5	40.2	12.0	19.3	30.3	42.8
	CR	13.9	22.0	29.3	42.5	14.4	22.9	32.3	45.2
	FLR	15.2	20.8	31.5	41.2	15.5	20.9	31.9	39.9
> 54	OLS	12.5	20.1	35.4	46.3	13.1	20.2	37.5	48.1
	CR	15.2	24.0	37.6	48.9	15.9	24.1	39.9	50.7
	FLR	19.8	28.7	41.8	50.0	21.0	24.7	42.2	49.2

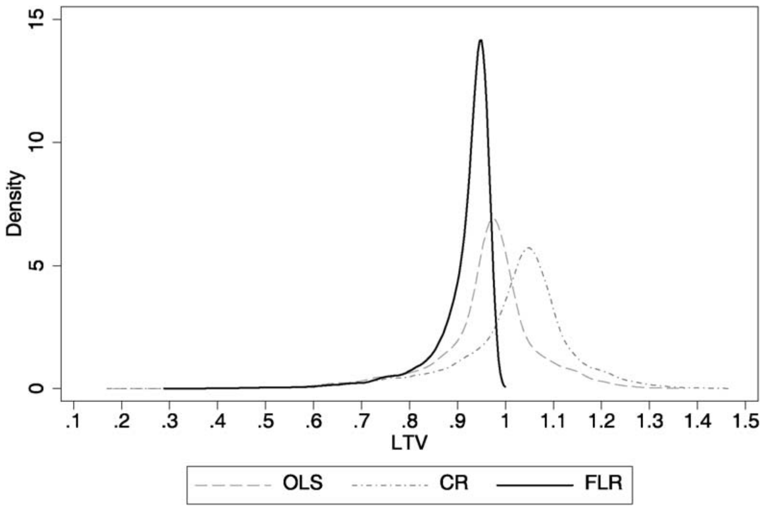

Fig. 4. Kernel density plots: simulated LTVs for zero tax penalty.

with LTVs greater than unity, considerably greater in the case of FLR. The FLR simulations, which are capped at unity, are necessarily lower.

On the other hand, the FLR estimated response exceeds the OLS estimated response for borrowers with moderate initial LTVs, and for whom the upper limit on debt gearing is unlikely to bind. While the OLS slope is restricted to be the same for all values of the independent variable, the FLR slope is not. This is clear from Fig. 2 where the central section of the FLR curve is steeper than the OLS line. And older borrowers in our sample are much more likely to lie on this section of the curve than are younger borrowers.

The impact of removal of the MID is much greater for POs than for FTOs and is again estimated to be much larger by the FLR method. Younger households have less wealth and are thus less able to reduce debt in response to a debt tax penalty. Moreover, most of the OLS and CR simulations of LTVs above unity are for younger borrowers. And POs have likely accumulated more wealth than FTOs.

Next we calculate the aggregate mortgage debt of borrowers both with and without interest deductibility implied by the three estimation techniques. Here we multiply the simulated LTVs with (without) deductibility for each household by its house value to get the simulated loan size (capped at the house value for FLR) with (without) deductibility. Then we compute the averages for each of the cells in Table 2 and multiply them by the relevant weights for each cell. Adding these products gives the average loan size with and without interest deductibility, respectively. The percentage decline in average loan size is obtained by dividing the difference between the averages with and without deductibility (times 100) by the average with full deductibility. This is also the percentage decline in aggregate mortgage debt due to the removal of the MID.

Two sets of weights are given in Table 5. Each set distributes homeowners into three age classes (< 35, 35–44, > 44), allowing for FTOs and POs as well as low and high tax brackets. The first set is based on our origination data in Table 2 and is derived directly from the counts in that table. These weights are necessarily skewed toward young, mobile households who do most loan origination. To compute the aggregate debt reduction for the economy, we want weights based on the home-owning population, not just new borrowers. The second set of weights is based on the British Household Panel Survey (BHPS) which is based on a representative sample of British households.[10]

In this case, a full 96 percent weight is applied to previous owners in the 40 percent tax bracket. Because most owners have been in their houses for many years (the average age of previous owners is 42.3 years in the BHPS sample versus 34.6 years in the new originators sample),

Table 5
Cell weights

	FTO		PO	
Age:	0.23–0.25	0.4	0.23–0.25	0.4
(a) CML cell weights				
< 35	0.336	0.034	0.155	0.061
35–44	0.057	0.017	0.100	0.077
> 44	0.026	0.007	0.083	0.048
(b) BHPS cell weights				
< 35	0.003	0.009	0.012	0.217
35–44	0.001	0.003	0.005	0.309
> 44	0.001	0.005	0.004	0.430

[10] We extracted a subsample of mortgage borrowers four years (1995–1998) of BHPS data which we treated as repeat cross sections, yielding a total of 11,313 observations. We used the BHPS rather than one of the cross sectional databases because it proved to be the only data set that could provide us with a sufficiently large number of observations on all of the variables needed to derive the cell weights: (1) whether the head of household was a previous owner; (2) date of purchase of dwelling; (3) price paid for dwelling; (4) value of outstanding mortgage debt; and (5) basic income of head of household.

Table 6
Percent decline in aggregate mortgage debt

	CML weights	BHPS weights
OLS	−20.9	−38.6
CR	−22.7	−40.6
FLR	−16.9	−32.3

incomes are higher in the BHPS sample. Further, the average LTV for the BHPS sample is only 0.61 versus the 0.78 in the CML sample. Recall that the tax calculation is based on income including an estimate of the income that would have been earned if the household were not a homeowner, i.e., the product of the mortgage interest rate and the difference between the current house value and the loan balance is added to other income.[11] This imputation pushes many more BHPS than CML households into the 40 percent tax bracket). In the absence of this imputation, only 50.2 percent of POs in the BHPS sample would be in the top tax bracket.

The aggregate percentage debt declines from removing the MID based on OLS, CR, and FLR are listed in Table 6 for the two sets of weights. Using the loan origination weights, the aggregate debt decline is 17 to 23 percent. Using the homeowner population weights, the responses rise to 32 to 41 percent. The latter, our CR estimate, is identical to the 41 percent simulated aggregate decline based on analysis of US data obtained by Follain and Melamed [7].[12] What we wish to emphasize, though, is the 20 percent lower FLR estimate of 32 percent. This lower response is attributable to capping the simulated LTVs with full MID at unity. As just shown, OLS and CR simulated values with the MID often exceeded unity. Thus the reductions from removal of the MID based upon these estimations are greater than those based on FLR. Because the household LTV response acts to offset the impact of removing the MID, a smaller response means a larger impact on housing demand and government tax revenue.

7. Conclusions

Given the non-taxation of income from owner-occupied housing, interest deductibility creates neutrality between debt and equity financing and renders the user cost of capital (through the WACC) independent of LTV choice. Removing, fully or partially, deductibility creates a debt tax penalty and makes the user cost a positive function of leverage. Thus the impact of changes in deductibility on housing choices, both tenure and quantity demanded, would depend on borrower sensitivity to the debt tax penalty. Estimating this sensitivity is the purpose of this paper.

The UK gradually removed the home mortgage interest deduction during the 1974–1999 period. During this period the penalty depended on whether the loan was above or below £30,000. Thus we estimate a two-equation model determining both the tax penalty based on the probability of a loan being above £30,000 and the LTV itself.

Explaining the LTV is complicated by the restraint that lenders place on the maximum value; 15 percent of our UK loans have an LTV of exactly 95 percent. Not only does this make the

[11] An estimate of the current house value in the BHPS is obtained by inflating the original purchase price by the house price deflator for the region is which the house is located. The deflator was based on the Nationwide UK regional price index available from http://www.nationwide.co.uk/hpi/historical.htm.

[12] Like Follain and Melamed, we have held house value constant. Removing the MID would, of course, reduce some combination of the price and quantity of housing demanded, leading to an even greater decline in mortgage debt. The larger is the leverage response to removal of the MID, the less will be the impact on housing demand.

explanation tricky, but linear extrapolation of borrower responses to full interest deductibility is likely to lead to predictions of many borrowers having LTVs exceeding what we know to be the lender maximum. We address these complications by considering three estimation methodologies: Ordinary Least Squares, Censored Regression (CR) and Fractional Logit Regression (FLR). Only the latter restricts predicted LTVs to be less than a maximum.

We analyze LTVs of 86,000 home purchasers with a mean LTV of 78 percent (based on a random sample of all new purchase loans). The simulated responses vary widely based upon the age of the borrower and whether they are first-time or previous owners. To illustrate, the FLR percentage declines for first-time owners are 7 and 16 for ages 25–34 and 45–54. For previous owners, the comparable age responses are 12 and 35. The simulated weighted-average percentage decline in LTV for previous owners in response to a shift from full interest deductibility to no deductibility ranges from 23 percent for Censored Regression to 17 percent for FLR. The much lower FLR response is due to the constraint against simulated LTVs exceeding unity when full interest deductibility exists.

These aggregate mortgage responses reflect our home buyer sample, in which first-time homeowners and younger buyers are far over represented relative to their distribution in the total homeowner population. Computations based on the overall population, i.e., with far greater weight given to older, wealthier, previous owners, yields greater percentage declines in aggregate mortgage debt. The Censored Regression decline is virtually identical to the economy-wide response simulated by Follain and Melamed [7] for the US. However, the FLR 32.4 percent decline is 20 percent less.

Acknowledgments

We thank Todd Sinai and the editor for their comments on an earlier version of this paper.

Appendix Table
Probit results (dependent variable = ceil)

	London/SE	Other regions
PO	−0.424	−0.193
	(−11.8)	(−8.3)
Income	0.048	0.082
	(12.8)	(18.3)
other income	0.027	0.073
	(7.5)	(22.8)
other income dummy	0.395	0.293
	(11.6)	(13.1)
age < 25	−0.361	−0.287
	(−2.8)	(−3.4)
age 25 to 34	0.946	0.484
	(11.6)	(7.2)
age 35 to 44	1.112	0.899
	(14.9)	(13.7)
age 45 to 54	0.373	0.664
	(4.8)	(10.1)
inc. ∗ (age < 25)	0.168	0.173
	(12.1)	(18.9)

(*continued on next page*)

Appendix Table (*continued*)

	London/SE	Other regions
inc. ∗ (age 25 to 34)	0.034	0.085
	(6.3)	(15.0)
inc. ∗ (age 35 to 44)	0.004	0.018
	(0.7)	(3.5)
inc. ∗ (age 45 to 54)	0.015	−0.012
	(2.9)	(−2.4)
PO ∗ (age < 25)	−0.052	0.206
	(−0.6)	(3.1)
PO ∗ (age 25 to 34)	0.084	0.148
	(1.6)	(4.5)
1996	0.001	0.034
	(0.0)	(1.8)
1997	0.053	0.130
	(1.8)	(6.9)
1998	0.059	0.118
	(1.8)	(5.9)
S.East	−0.115	–
	(−4.8)	–
Y.Hum	–	0.244
	–	(8.5)
E.Mids	–	0.263
	–	(8.9)
N.West	–	0.305
	–	(10.8)
Scot.	–	0.253
	–	(8.3)
W.Mid	–	0.430
	–	(14.5)
S.West	–	0.555
	–	(19.2)
E.Ang.	–	0.411
	–	(12.0)
Const.	−0.405	−1.802
	(−5.7)	(−27.9)
N	33,541	53,079
LL	−8568	−21,700
Chi^2	5270	14,000
AIC	17,176	43,418

Note. Figures in parentheses are t-ratios.

References

[1] T. Amemiya, Regression analysis when the dependent variable is truncated normal, Econometrica 41 (1973) 997–1016.

[2] H. Bester, The role of collateral in credit markets with imperfect information, European Economic Review 31 (1987) 887–899.

[3] J.K. Brueckner, Consumption and investment motives and the portfolio choices of homeowners, Journal of Real Estate Finance and Economics 15 (1997) 159–180.

[4] R.M. Dunsky, J.R. Follain, Tax-induced portfolio reshuffling: The case of the mortgage interest deduction, Real Estate Economics 28 (2000) 683–718.

[5] J.R. Follain, R.M. Dunsky, The demand for mortgage debt and the income tax, Journal of Housing Research 8 (1997) 155–199.

[6] J.R. Follain, D.C. Ling, The federal tax subsidy to housing and the reduced value of the mortgage interest deduction, National Tax Journal 44 (1991) 253–266.
[7] J.R. Follain, L.S. Melamed, The false messiah of tax policy: What elimination of the home mortgage interest deduction promises and a careful look at what it delivers, Journal of Housing Research 9 (1998) 179–199.
[8] W.H. Greene, Econometric Analysis, Macmillan, New York, 1993.
[9] P.H. Hendershott, J. Slemrod, Taxes and the user cost of capital for owner-occupied housing, AREUEA Journal 11 (1983) 375–393.
[10] P.H. Hendershott, J.R. Follain, D.C. Ling, Effects on real estate, in: J. Pechman (Ed.), Tax Reform and the US Economy, The Brookings Institution, Washington, DC, 1987, pp. 71–94.
[11] P.H. Hendershott, W.C. LaFayette, D.R. Haurin, Debt usage and mortgage choice: The FHA-conventional decision, Journal of Urban Economics 41 (1997) 202–217.
[12] P.H. Hendershott, G. Pryce, M. White, Household leverage and the deductibility of home mortgage interest: Evidence from UK house purchasers, Journal of Housing Research 14 (2003) 49–82.
[13] D.C. Ling, G.A. McGill, Evidence on the demand for mortgage debt by owner occupiers, Journal of Urban Economics 44 (1998) 391–414.
[14] F. Modigliani, M. Miller, Cost of capital, corporation finance and the theory of investment, American Economic Review 48 (1958) 261–297.
[15] C. Mooney, R. Duval, Bootstrapping: A Nonparametric Approach to Statistical Inference, Sage, Newbury Park, CA, 1993.
[16] L.E. Papke, J.M. Wooldridge, Econometric methods for fractional response variables with an application to 401(k) plan participation rates, Journal of Applied Econometrics 11 (1996) 619–632.
[17] J. Stiglitz, A. Weiss, Credit rationing in markets with imperfect information, American Economic Review 71 (1981) 393–411.
[18] J. Tobin, Estimation of relationships for limited dependent variables, Econometrica 26 (1958) 24–36.
[19] S.E. Woodward, Weicher J.C., Goring the wrong ox: A defense of the mortgage interest deduction, National Tax Journal 42 (1989) 301–313.
[20] J.M. Wooldridge, Econometric Analysis of Cross Section and Panel Data, MIT Press, Cambridge and London, 2002.