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Bram De Rock
Mariia Kovaleva
Tom Potoms

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Housing wealth, marital stability and labor supply: an intertemporal analysis

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Housing Wealth, Marital Stability and Labor Supply: an Intertemporal Analysis*

Bram De Rock[†] Mariia Kovaleva[‡] Tom Potoms[§]

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Abstract

We study how house price shocks affect marital stability and household labor supply. We address this question using a dynamic collective household model with limited commitment. We find that positive house price shocks increase the divorce rate, and that leverage ratios such as loan-to-income (LTI) and loan-to-value (LTV) determine the transmission of house price shocks on divorce. Given its importance, we then analyze a tightening of the credit market through the LTI-limit. We show that neglecting the divorce and intra-household bargaining channels significantly biases the individual welfare effects of such policies.

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†ECARES, Université Libre de Bruxelles and KU Leuven. Avenue F. d. Roosevelt 50, 1050 Brussels, Belgium. E-mail: bram.de.rock@ulb.be. Bram De Rock gratefully acknowledges FWO and FNRS for their financial support.

‡ECARES, Université Libre de Bruxelles.Avenue F. d. Roosevelt 50, 1050 Brussels, Belgium. E-mail: Mariia.Kovaleva@ulb.be. Mariia Kovaleva gratefully acknowledges FNRS for their financial support

§University of Sussex. Jubilee Building 9SL, Falmer Brighton, BN1 9SN United Kingdom. E-mail: T.Potoms@sussex.ac.uk

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1 Introduction

Across advanced economies, rising house prices and mounting household debt have prompted a reassessment of how housing wealth fluctuations affect financial decision-making of households. In the United Kingdom—the focus of our study—house prices have increased by roughly 43 percent over the past decade (ONS, 2019), while a challenging global economic outlook further complicates households' ability to service their debt (Bank of England, 2022)). With housing wealth comprising nearly 50 percent of total household assets in OECD countries, even modest price changes can tighten borrowing conditions and reshape behavior.[1]

These macroeconomic pressures also influence household-level decisions, highlighting the importance of the relationship between marriage and homeownership, which plays a dual role in households. On the one hand, marriage enables couples to combine income sources and mitigate borrowing constraints. On the other, housing wealth serves as a form of insurance to spouses against divorce risk, helping couples with intra-household specialization.[2]

Structural interpretation of the data patterns. In this paper, we therefore examine how increases in house prices impact marital stability, female labor supply, and individual welfare. Our study contributes in three important ways. First, we provide robust empirical evidence that links house price shocks to marital instability and changes in female labor supply, with these effects varying systematically by household leverage measures such as debt-to-primary earner's income (pLTI), and loan-to-value (LTV) ratios. Second, we develop a dynamic structural model of house-

[1] See the 2019 OECD report on housing wealth by Causa et al. (2019).

[2] See Lafortune and Low (2023) for evidence that homeownership can buffer divorce risk, particularly for the household member specializing in domestic production.

hold behavior that integrates borrowing constraints, intra-household bargaining, and divorce dynamics—allowing us to disentangle the mechanisms through which house price shocks affect marital surplus and labor supply decisions. Third, we leverage our structural model to conduct counterfactual policy experiments to show that neglecting intra-household dynamics in the evaluation of credit market policies can lead to significant biases in the estimated welfare effects.

To that end, we proceed in two steps. We start by presenting reduced form evidence on how the leverage ratios LTV and pLTI play a role in the transmission of house price shocks in divorce outcomes of households. For this we use county-level house price data from the UK to construct a measure of house price shocks as the accumulated deviations from their trend in an approach similar to Disney et al. (2010) and Rainer and Smith (2010). Merging these with panel data of UK households from the British Household Panel Survey (BHPS), we show how households with a high pLTI ratio experience an increase in the probability of divorce in response to a positive house price shock, whilst households with high LTV ratios experience a decrease in the divorce probability.

In a second step, we build a dynamic structural model of the household behavior in which the intra-household decision making process is characterized by limited commitment (Mazzocco, 2007; Voena, 2015; Chiappori and Mazzocco, 2017; Lise and Yamada, 2019). This implies that spouses may revise their bargaining power whenever their participation constraint binds—making divorce an optimal choice if a spouse's outside option becomes sufficiently attractive. Borrowing constraints arise from the more binding of an LTV-limit and a debt-to-income (LTI) limit, the latter endogenously linked to female labor supply. Households jointly determine housing, labor, and consumption choices in each period. To capture variations in household outside options, the model incorporates shocks to match quality, productivity, and house prices. We further exploit quasi-experimental variation from the *White v. White* case in England—which shifted the wealth distribution upon divorce—to identify the initial intra-household allocation of resources and to validate our structural estimates.

Findings from counterfactual experiments. We conduct two sets of experiments with our structural model. We first consider an unexpected transitory 10% surge in house prices. We find that the shock reduces homeownership rates, lowers married women's employment, and increases divorce rates. The impact is particularly pronounced among younger households.

The key mechanisms driving these outcomes are intra-household bargaining and borrowing constraints, which exert distinct effects based on housing status. For households that transition from homeownership to renting as a result of the shock, the easing of leverage-based borrowing constraints reduces the necessity for women to work. Moreover, diminished expectations of future housing wealth tighten marital participation constraints. This tightening elevates divorce risk and shifts bargaining power toward wives, who subsequently increase leisure even further. Among homeowners, increased housing wealth generates both wealth effects and bargaining shifts, similarly reducing married women's employment. Additionally, high indebtedness—as captured by high debt-to-primary income (pLTI) ratios—further amplifies divorce risk following the shock, especially when adverse match quality shocks exacerbate bargaining tensions at later life stages. In contrast, higher loan-to-value (LTV) ratios mitigate this risk by relaxing financial constraints.

In a second experiment, we explore the long-run effects of tighter credit conditions by imposing a stricter loan-to-income (LTI) limit—a macroprudential measure aimed at curbing excessive borrowing and influencing marital stability. We simulate responses under three scenarios: (i) a limited commitment framework, where spouses can adjust their bargaining weights; (ii) a scenario with fixed intra-household bargaining but divorce allowed; and (iii) a scenario with fixed bargaining and no divorce.

Intra-household adjustments modify the impact of tighter credit on private consumption, female employment, homeownership, and individual welfare. Under the limited commitment framework, homeownership rates decline the most, and the employment rate exhibits the smallest increase. Tighter credit conditions enhance female bargaining power, which dampens the employment response and boosts private consumption—particularly for households transitioning from ownership to renting. In contrast, when bargaining power is fixed, households adjust primarily by increasing savings and altering labor supply more significantly. We find that neglecting intra-household bargaining dynamics and divorce leads to significant biases in assessing the individual welfare impacts of credit market policies. This bias arises because failing to account for these mechanisms overlooks the reallocation of resources—particularly adjustments in leisure and private consumption for married women—especially when households transition from homeownership to renting.

Relation to the literature. Our paper touches upon several strands of literature. First, we contribute to those papers that study how house price changes affect individual outcomes within the household. While much of the existing research, both empirical and structural, focuses on consumption (Campbell and Cocco, 2007; Disney et al., 2010; Mian et al., 2013; Berger et al., 2018; Etheridge, 2019; Guren et al., 2020; Kaplan et al., 2020), fewer studies examine the impact on household labor supply (Disney and Gathergood, 2018; Low and Sánchez-Marcos, 2024). Importantly, these studies either adopt a fully empirical approach or rely on unitary household models, which overlook intra-household bargaining and endogenous divorce dynamics. Our limited commitment household model addresses this gap by exploring the interaction between labor supply, house price shocks, and marital stability.

Second, we contribute to the broad research on household decisions under divorce risk (Fernández and Wong, 2014; Voena, 2015; Fischer and Khorunzhina, 2019). We place particular emphasis on the relatively sparse work examining marital instability and changes in housing wealth. Most existing studies, such as Rainer and Smith (2010) and Farnham et al. (2011), are fully empirical and do not account for borrowing constraints as a channel through which house price shocks are transmitted. We build on recent work highlighting the role of leverage ratios, such as LTV and LTI, in the transmission of house price shocks to housing investment (Cloyne et al., 2019; Crossley et al., 2024; Ahlfeldt et al., 2024). Our quantitative analysis complements this literature by showing how endogenous leverage constraints and intra-household bargaining shape both the direction and magnitude of house price effects.

Finally, our policy counterfactual exercise focuses on macroprudential policy, connecting this paper to works by Attanasio et al. (2012), Pizzinelli (2018) and Bartscher (2023). Unlike these papers, we quantify how regulatory changes affect individual welfare by incorporating intra-household bargaining and divorce dynamics.[3]

Outline. The rest of the paper is organized as follows. Section 2 describes our data. It also presents the reduced-form evidence that motivates and disciplines our structural model. Section 3 presents the dynamic model. Section 4 discusses identification and the estimation results of our model. Section 5 analyzes the responses to a house price shock and the importance of leverage ratios in the transmission of such shocks to marital (in-)stability. Section 6 presents the policy counterfactual analysis. Finally, Section 7 concludes. Additional details and results are all contained in the Online Appendix.

2 Empirical analysis on labor supply, housing wealth, and divorce

In this section, we document two complementary strands of reduced-form evidence: (i) how exogenous house price shocks interact with borrowing constraints to affect divorce probabilities, and (ii) how the post-divorce division of housing wealth influences household decision-making. These findings motivate the structural framework developed in Section 3. We start with presenting the data sources that we will use throughout the paper.

2.1 Data sources

For the analysis in this paper we will need information on marital transitions, labor supply, wages, earnings, homeownership, house prices, and further information on

[3]The latter distinguishes our analysis from existing studies on the welfare effects of house price shocks (see, for example, Fagereng et al. (2024)) which center on redistribution due to rising asset valuations and declining discount rates.

households' balance sheets. To obtain this information we will work with several data sources throughout.

The first dataset is the British Household Panel Survey (BHPS). The initial wave of the BHPS contained approximately 5,500 households, which sums to about 10,000 respondents. In addition to this basic sample, representative for the national population of the UK, there were booster samples between 1997-2001 consisting of lower-income individuals and from 1999 there was a boost of respondents from both Scotland and Wales. Our main estimating sample from the BHPS will consist of respondents between 25-60 years old living in England or Scotland.[4] The sample range is restricted to 1992-2005, given that some controls at the country-level are not available in 1991 and in 2006 there was the introduction of the Scottish Family Law, which could act as a confounder to the analysis.[5] To increase power on divorce rate measures, we will additionally use data on crude divorce rates from the Office for National Statistics (ONS), 'Vital statistics: Population and Health Reference tables' (years 1990-2005).

Next, we need information on house prices. For our reduced-form analysis we will make use of local authority (LA)-level yearly prices from semi-detached houses (years 1995-2005), obtained via the Halifax Bank of Scotland, which is a major mortgage lender in the UK. Given that we have access to the LA of our respondents within the BHPS, we can then easily merge the Halifax house price indices at the county level with our main estimation sample from the BHPS.

Finally, we will use data from the UK House Price Index (UKHPI) 1968-2008 to obtain good estimates of the UK-wide house price process and variance of house price innovations for the structural estimation.

2.2 House price shocks, divorce and leverage ratios

We begin by presenting reduced-form evidence on how house price shocks affect household divorce likelihood and the role of leverage-based constraints in transmitting housing price volatility to divorce outcomes.

[4] Similar to Piazzalunga (2017), we do not consider Wales in the empirical analysis.

[5] We follow the convention to refer to England, Wales and Scotland as 'countries', whereas the metropolitan and government regions will simply be referred to as 'regions'.

The main challenge in this analysis is obtaining an exogenous measure of house price shocks. Even though the BHPS includes self-reported housing values, these data are problematic: they are highly endogenous and only available for homeowners. Consequently, we will follow Disney et al. (2010) and Rainer and Smith (2010) to construct a more plausibly exogenous measure of a house price shock. Specifically, we employ Halifax county-level house price data to estimate an AR(2) process for (the log of) LA-level house prices, and accumulate the residuals from this regression as a measure for the house price shock.[6]

We then estimate an individual-level regression where the dependent variable is divorce likelihood on the constructed house price shock measure at local authority level l in which the household h lives, $HPshock_{l,t}$.[7] We interacted $HPshock_{l,t}$ with the standard measures of leverage ratios that are relevant to indicate borrowing constraints for households. More specifically, we include the lagged value of primary loan-to-income ratio $pLTI_{h,t-1}$, defined as the amount of outstanding mortgage debt divided by the primary earner's income, as well as the lagged value of the loan-to-value ratio, $LTV_{h,t-1}$.

Table 1 presents our estimates of individual divorce likelihood as a function of house price shocks and their interactions with leverage measures. In column 1 we present the results for all respondents regardless of housing tenure, whilst column 2 restricts the sample to homeowners. Overall, our findings indicate that house price shocks affect divorce likelihood in ways that depend on households' borrowing constraints. Specifically, a positive house price shock combined with a high primary debt-to-income (pLTI) ratio increases the risk of divorce, likely because tighter borrowing constraints limit these households' ability to adjust their housing consumption. In contrast, a higher loan-to-value (LTV) ratio appears to mitigate marital instability by effectively relaxing borrowing constraints (Crossley et al., 2024). These results are suggestive for the fact that leverage ratios (borrowing capacity) affect the transmission of house price shocks on marital surplus.

[6]We refer to Online Appendix A for more details.

[7]Divorce is measured in the BHPS as a dummy variable that equals 1 if marital status changes from married to divorced between two waves, and 0 otherwise.

Table 1: *House price shocks and probability of divorce.*

	Prob. Divorce$_{h,t}$	*Prob. Divorce*$_{h,t}$
$HPshock_{l,t} \times pLTI_{h,t-1}$	0.011***	0.011***
	(0.0041)	(0.0042)
$HPshock_{l,t} \times LTV_{h,t-1}$	-0.055	-0.049
	(0.052)	(0.055)
$pLTI_{h,t-1}$	0.0005	0.0004
	(0.001)	(0.001)
$LTV_{h,t-1}$	-0.0073**	-0.0042
	(0.003)	(0.003)
$HPshock_{l,t}$	-0.009	-0.012
	(0.017)	(0.020)
Sample	all	homeowners
Observations	30,401	25,819
R-squared	0.0084	0.0078
Demographic controls	✓	✓
Year	✓	✓
Region	✓	✓

Notes: estimation on the BHPS sample, respondents in the age range of 25-60 years old during the period 1995-2005. Standard errors clustered at the individual level, *** p<0.01, ** p<0.05, * p<0.1. Demographic controls include the duration and duration squared of the couple, age and age squared of the respondent, the education level of the respondent defined as their highest qualification received, number of young children ($\leq$ 15 age old) and the household's non-labor income. Non-labor income is deflated by the CPI with 2014 as reference year. Cross-sectional weights have been used.

2.3 Housing wealth division post-divorce, labor supply, and marital instability

Another key determinant in how housing wealth affects intra-household decisions is how it is split upon divorce. Given that we use data from the UK we can exploit the so-called *White v. White* case, which changed the manner in which assets are split post-divorce. We can use this policy change to assess how increased access to wealth post-divorce affects household decisions and (crude) divorce rates. This is important for our structural estimation, as these reduced-form estimates will be used to identify our dynamic model.

Background. The *White v. White* decision is considered a decisive change in the pattern of property and asset division upon divorce for English households. As discussed in Chiappori and Mazzocco (2017), this case essentially meant a shift from a separate ownership of assets to an equal division. In addition, due to a different divorce legislation in Scotland, the reform only affected English and not Scottish households, which allows us to treat them as a control group in our analysis.

In terms of background, Mr. and Mrs. White were business partners of a farming business in Somerset. At the time when the court case regarding their divorce came up, their combined net wealth was estimated at approximately 4.5 million GBP. Initially, Mrs. White was awarded a sum of 980,000 GBP, to which she appealed. The Court of Appeal then granted her 1.5 million GBP, using a 'yardstick of equality'. This decision was then confirmed by a ruling from the House of Lords in October 2000, where Lord Justice Nicholls in particular specified that, when a couple starts with a small amount of assets, which then grow considerably over the course of the marriage, both spouses, including the wife, should expect to receive half of that accumulated wealth, even if she has "never or rarely worked outside the home" (Stowe, 2009). Another argument was made by Lord Justice Thorpe, who argued that typically the wife "sacrifices her potential to generate assets by taking on the domestic commitment to her husband and her children." All this implied that, ever since this case, "the 50/50 split is, more often than not, a given" (Stowe, 2009).

Though most of the practical applications of this case would involve 'big money' cases, it is widely acknowledged that the *White v. White* case had a much broader impact on the post-divorce division of assets (Smith, 2003). In addition, the case was widely promoted with broad media coverage, thereby informing potential married couples about their likely more egalitarian shares of joint assets upon divorce (Piazzalunga, 2017).[8]

[8] More institutional background on the asset division in English divorce law can be found in Online Appendix B.

Effects on labor supply and divorce. To study the effect of the *White v. White* case on household decisions, we employ a traditional Difference-in-Difference (DiD) regression, with the treatment group defined as married women living in England and the control group married women living in Scotland. In Online Appendix C, we included the results of several extra exercises to confirm the validity of our DiD approach.

We run a regression of the following form:

$$Hours_{i,c,t} = \phi Post \times Treated_{c,t} + \gamma \mathbf{X'}_{i,c,t} + \sum_t f_t + \sum_r f_r + \epsilon_{i,c,t}, \tag{1}$$

where *Hours* denotes the number of working hours for married women i, living in country c, in year t.[9] The variable $Treated_{c,t}$ is an indicator for being in the treatment group, that is, $Treated_{c,t} = 1$ if the respondent is living in $c =$ England in year t. The dummy variable *Post* equals one if the year $\geq$ 2000. We always include a full set of region and time dummies, f_r and f_t, as well as demographic controls $\mathbf{X'}_{i,c,t}$ at the household and individual level.[10] Furthermore, we control for the country-level female unemployment rate.

For marital instability, we turn to crude divorce rates to cover the aggregate effects of the reform.[11] We use a similar set-up for the (crude) divorce rate, except of course for the control variables and the inclusion of a linear time trend τ:

$$Divorce_{c,t} = \tilde{\phi} Post \times Treated_{c,t} + \sum_c f_c + \sum_t f_t + \sum_c f_c \times \tau + \epsilon_{c,t}. \tag{2}$$

[9]There are several measures for working hours available in the BHPS. The three most common definitions of hours worked involve (i) contractual hours worked, (ii) contractual hours + hours of paid overtime and (iii) contractual hours + total overtime hours. We also experimented with several definitions of being employed in the BHPS, e.g., based on whether or not the respondent has done paid work in the week leading up to the interview, an alternative measure based on self-reported employment status and finally one based on reported hours worked. All these different measures did not change the conclusions presented in the main text.

[10]Specifically, we include the age and age squared of husband and wife, the education level of husband and wife both defined as their highest qualification received, number of young children ($\leq$ 15 age old) and the household's non-labor income.

[11]The household-level divorce outcomes are likely to be under-powered, therefore we follow Piazzalunga (2017) to look at crude divorce rates as the dependent variable.

The results in Table 2 show that there is evidence for an average reduction of married women's working hours by about 2.2 hours per week. We also ran a similar specification as in (1) with employment status as outcome variable, which did not yield any significant results. This is completely in line with earlier findings by studies on the *White v. White* case (Kapan, 2008; Piazzalunga, 2017). In terms of the crude divorce rates, there is an uptick of about 1.2 divorces per 1,000 persons.[12]

Table 2: *White v. White: labor supply of married women and crude divorce rates.*

	$Hours_{i,c,t}$	$Divorce_{c,t}$
$Post \times Treated_{c,t}$	-2.23**	1.22**
	(0.925)	(0.49)
Data source	*BHPS*	*Vit. Stats*
Observations	26,643	30
R-squared	0.20	0.98
Demographic (incl. spouse) controls	✓	
Year	✓	✓
Region	✓	
Local female unemployment rate	✓	
Country		✓
Country $\times$ τ		✓

Notes: estimation on the BHPS sample, married women in the age range of 25-60 years old during the period 1992-2005. Standard errors clustered at the individual level, *** p<0.01, ** p<0.05, * p<0.1. Demographic controls include age and age squared of the respondent, the age and age squared of the husband, the education level of the respondent and the education level of the husband both defined as their highest qualification received, number of young children ($\leq$ 15 age old) and the household's non-labor income. Non-labor income is deflated by the CPI with 2014 as reference year. Cross-sectional weights have been used. The second regression is based on the Vital Statistics data for the period 1990-2005.

Although our reduced-form analyses robustly document the empirical link between house price shocks, borrowing constraints, and marital stability, they do not disentangle the underlying mechanisms—namely, the interaction of intra-household bargaining, credit constraints, and selection into homeownership. In Section 3, we

[12]We also experimented with the addition of a country-specific quadratic time trend, following Friedberg (1998) and Piazzalunga (2017). This did not affect the point estimate much, only lowering the precision.

develop a dynamic structural model that formalizes these channels and quantifies their welfare implications.

3 Structural model

In this section, we describe our lifecycle model of household behavior. Married individuals make *Pareto optimal* decisions about consumption, housing demand and labor supply under *limited commitment*. This implies that the household maximizes a weighted sum of the spouses' utility where the weights reflect each spouse's relative bargaining power. In addition, spouses cannot commit to future allocations. Hence, the bargaining weights can be revised whenever one of the spouses becomes better off outside of marriage.

3.1 General set-up

Time is modeled in discrete periods, indexed by t. The lifecycle is divided into two phases: a working phase ($t \leq T_r$) and a retirement phase ($T_r < t \leq T_d$). During the working phase, married individuals jointly decide on private consumption, labor supply, and housing demand. At the end of this phase, they retire and cease making labor supply decisions. Additionally, married individuals may unilaterally choose to divorce, after which they remain single for the rest of their lifecycle.

In what follows, we outline the key components of the model and formulate the household's optimization problem. The main text focuses on the optimization problem for married individuals during the working phase. For further details on the retirement phase—where income becomes exogenous (i.e., through pensions)—and the analogous set-up for singles (which is similar to the married case except for the absence of joint decision-making and economies of scale), please refer to Online Appendix D.

3.2 Preferences

We study opposite-sex couples where variables referring to the husband are indicated by M and for the wife with F. In each period t, couples make decisions over private consumption $c_{M,t}, c_{F,t}$, female labor supply (leisure), $n_{F,t}$ ($l_{F,t} = 1 - n_{F,t}$), and housing, $H_t \in \{1, 2, 3\}$. We follow Yang (2009) and interpret H_t as a service flow of housing. Here, $H = 1$ indicates the decision to rent, and for ease of notation we think of $H = 2$ as services obtained from flat ownership and $H = 3$ as service flows obtained from owning a house.[13]

With respect to private consumption, we follow Voena (2015) and assume that households are characterized by economies of scale. In particular:

$$x_t = F\left(c_{M,t}, c_{F,t}\right) = \left[(c_{M,t})^{\rho} + (c_{F,t})^{\rho}\right]^{\frac{1}{\rho}},$$

where x denotes total private expenditures of the household. If $\rho \geq 1$, then this functional form implies that the spouses can consume more jointly, compared to what they would if they lived separately.

The intra-period utility of a married individual ($g \in \{M, F\}$) takes on the following form:

$$u_g\left(c_{g,t}, l_{g,t}, H_t; \theta_{g,t}\right) = \frac{c_{g,t}^{1-\sigma_g}}{1-\sigma_g} + \omega_{g,mar}^{l} \frac{l_{g,t}^{1-\psi_g}}{1-\psi_g} + \omega_{g,mar}^{H} \mathbf{1}\left[H_t > 1\right] + \theta_{g,t}.$$

Some remarks are in order. First, in line with the data, we assume that male labor supply is more stable. We therefore consider the case where men work full time until retirement, meaning that we can ignore the utility from leisure for them.[14] Women's leisure is an explicit choice variable in the model, where $\omega_{F,mar}^{l}$ is a weight attached to her utility of leisure, while ψ_F determines her Frisch elasticity of labor supply. We chose not to include household work as a choice variable, not only for computational

[13]In the remainder of the paper we will for simplicity refer to $H = 2$ as 'flat ownership' and $H = 3$ as 'house ownership'.

[14]We do allow for some variation in hours worked for married men over the lifecycle through (exogenous) job displacement shocks, see infra.

reasons, but also given that for identification purposes, there is no empirical evidence that parents have substituted market work for domestic work after the *White v. White* reform (see Online Appendix C for more details).

Second, housing is a public good and yields a marginal utility equal to $\omega^{H}_{g,mar}$, to be more specific, we have:

$$\omega^{H}_{g,mar} = \omega^{H,0}_{g,mar} + \omega^{H,1}_{g,mar} \times (H - 1). \tag{3}$$

Where $\omega^{H,0}_{g,mar}$ is the marginal utility for homeownership, whilst $\omega^{H,1}_{g,mar}$ gives the relative preference for the size of housing. Note that we allow for the housing preferences to depend on both gender as well as marital status. This flexibility is a parsimonious way to capture differences in utility from housing dependent on different household composition by marital status.

Finally, each spouse receives an individual-specific match quality shock, $\theta_{g,t}$, capturing the non-material benefits of marriage. We assume that these shocks follow a unit root process:

$$\theta_{g,t} = \theta_{g,t-1} + \epsilon^{\theta}_{g,t},$$

with $\epsilon^{\theta}_{g,t} \sim N\left(0, \sigma^2_{g,\theta}\right)$.

3.3 Sources of uncertainty

Besides match quality ('love'), individuals face the following sources of uncertainty.

House and flat prices. We denote by $\tilde{p}_t\left(H\right)$ the price of housing services at level H, with the following specification:

$$\tilde{p}_t\left(H\right) = \begin{cases} p_t \text{ if } H = 3, \\ \kappa_F p_t \text{ if } H = 2, \\ \kappa_R p_t \text{ if } H = 1. \end{cases}$$

We follow Attanasio et al. (2012), and assume that house prices p_t follow an AR(1) process, with a deterministic trend that reflects an upward drift. Specifically,

$$\ln p_t = a_0 + a_1 t + \rho_H \ln p_{t-1} + \varepsilon_{H,t},\ \ \varepsilon_{H,t} \sim N\left(-\frac{\sigma_H^2}{2}, \sigma_H^2\right). \tag{4}$$

The prices of a flat and the rental prices are given by a constant fraction, resp. κ_F and κ_R, of house prices. In that sense, one can interpret (4) as specifying a relative house over flat price process and similar for renting.

Unemployment shock for men. As indicated above, men work full time over the entire working phase of the lifecycle ($t \leq T_r$). However, we assume that in each period they face the risk of becoming (involuntarily) unemployed with an associated probability π^u. Their employment status can therefore be indicated by a variable $e_{M,t} \in \{0, 1\}$, with $e_{M,t} = 1$ indicating employment and $e_{M,t} = 0$ indicating unemployment. When unemployed they receive an unemployment benefit given by b^u. To simplify the analysis, we assume that the probability of becoming unemployed in each period is independent of the employment status in the previous period. We leave a full analysis allowing for unemployment persistence and the impact of subsequent human capital depreciation and household income for future research.

Earning and wage process. The earning process of men is given by

$$\ln y_{M,t} = \alpha_M(t) + \nu_{M,t},$$

and the wage process of women by

$$\ln w_{F,t} = \alpha_F(t) + \nu_{F,t}.$$

Both men and women face a concave lifecycle profile in earnings (or wages) :

$$\alpha_g(t) = \alpha_{g,1} t + \alpha_{g,2} t^2,\ \ g \in \{M, F\}.$$

The permanent shocks, $\nu_{g,t}$ reflect shocks in productivity, health, etc. Following Blundell et al. (2008), we assume that these follow a random walk:

$$\nu_{g,t} = \nu_{g,t-1} + \varepsilon_{g,t}, \ g \in \{M, F\}.$$

Finally, we allow for correlation in spouses' permanent shocks. In particular, $\varepsilon_t = (\varepsilon_{M,t}, \varepsilon_{F,t}) \sim N(\mu_\varepsilon, \Sigma_\varepsilon)$, with

$$\mu_\varepsilon = \left(-\frac{\sigma^2_{\varepsilon_M}}{2}, -\frac{\sigma^2_{\varepsilon_F}}{2}\right) \qquad \text{and} \qquad \Sigma_\varepsilon = \begin{pmatrix} \sigma^2_{\varepsilon_M} & \sigma_{\varepsilon_M,\varepsilon_F} \\ \sigma_{\varepsilon_M,\varepsilon_F} & \sigma^2_{\varepsilon_F} \end{pmatrix}.$$

The correlation in spouses' productivity shocks captures the fact that in reality, through assortative matching patterns in the marriage market, intra-household income shocks are correlated, which can have serious implications for intra-household inequality and the evolution of income inequality (Fernández and Rogerson, 2001; Lise and Seitz, 2011; Eika et al., 2019; Chiappori et al., 2020).

3.4 Budget and borrowing constraints

Budget constraint. The household faces in each period t the following budget constraint:

$$\begin{aligned} & A_t + x_t + \tilde{p}_t(H_t) H_t + \Omega(H_t, H_{t-1}) \\ = & (1+R) A_{t-1} + Inc_{M,t} + w_{F,t} n_{F,t} + y^{nl}_t - \delta^{mar}(t) \mathbf{1}[n_{F,t} > 0] \\ & + \tilde{p}_t(H_{t-1}) H_{t-1} \mathbf{1}[H_{t-1} > 1]. \end{aligned} \tag{5}$$

y^{nl}_t denotes the household's nonlabor income and household resources are also augmented by returns from savings The income of the husband, $Inc_{M,t}$ is given by either his labor income (in case he works $e_{M,t} = 1$), or he receives an unemployment benefit b^u in case he is (involuntarily) unemployed ($e_{M,t} = 0$). In sum,

$$Inc_{M,t} = y_{M,t} e_{M,t} + b^u (1 - e_{M,t}).$$

In terms of expenditures, we incorporate a cost of working for women, which, similar to Borella et al. (2018), we express in monetary terms by $\delta^{mar}(t) = \delta_1^{mar} t + \delta_2^{mar} t^2$. Second, note that wealth effects from housing stock are captured by the change in housing values and housing status. We assume households face a transaction cost $\Omega(H_t, H_{t-1})$ in case $H_t \neq H_{t-1}$, proportional to the housing value and asymmetric for selling and buying. Also note that when the household owns their property ($\mathbf{1}[H_{t-1} > 1]$), then they can experience a wealth effect if their property appreciates in value. Finally, and following Bajari et al. (2013), we should highlight that the interest rate, R, for assets is realistically allowed to vary depending on the sign of household savings. In particular, we assume that $R = R^-$ in case $A < 0$, which implies that the household is a net debtor, while $R = R^+$ in case the household is a net creditor, i.e., when $A \geq 0$. In addition, $R^- > R^+$. This captures the idea that interest rates are generally higher for mortgages than for saving accounts.

Borrowing constraint. An important aspect of our model is that households face leverage-based borrowing constraints, i.e., the amount of debt they can accumulate is determined by their housing status. More specifically, following Attanasio et al. (2012) and Pizzinelli (2018), we assume:

$$
\begin{aligned}
Debt_t &\leq \max\{Debt_{t-1}, \min\{LTV_t, LTI_t\}\}, \qquad (6)\\
LTV_t &= \lambda^H \tilde{p}_t(H_t),\\
LTI_t &= \lambda^Y (Inc_{M,t} + w_{F,t} n_{F,t}).
\end{aligned}
$$

The policy parameter λ^H captures how much the household can borrow as a fraction of the (fluctuating) housing values (i.e., the loan-to-value), which essentially translates to a downpayment restriction. The LTI-limit reflects that the household can only borrow up to a certain multiple (λ^Y) of household income. The debt in period t can only increase compared to the previous period, in case the outstanding debt is still meeting both leverage constraints.

We follow Bottazzi et al. (2007) and Pizzinelli (2018) in assuming that net liquid savings constitutes a single continuous variable and hence we do not allow for a separate choice of mortgage contract and deposits. Obviously, in reality households hold both positive liquid assets and mortgage debt, but decoupling these would significantly add to the computational complexity of our model. Interestingly, our focus on illiquid assets such as housing and neglecting the diversity in the household balance sheet, can be justified by a renewed recent focus on so-called "wealthy hand-to-mouth"-type consumers. These are consumers with large illiquid assets (such as housing), but relatively low amounts of liquid assets. This has direct consequences for example in the context of macroeconomic stabilization policies, see Kaplan et al. (2014), Kaplan and Violante (2014), Kaplan et al. (2018).[15]

3.5 Household's optimization problem

We now turn to a description of the households' optimization problem, which includes our specification of limited commitment.

Household's value function. Let $\mathbf{S}_t = \left(A_{t-1}, H_{t-1}, p_t, w_{F,t}, y_{M,t}, e_{M,t}, y_t^{nl}, \theta_{M,t}, \theta_{F,t}, \tilde{\mu}_{M,t}, \tilde{\mu}_{f,t}\right)$ denote all the relevant state variables relevant for the household's decision problem in period $t \leq T_r$. This comprises the asset (debt) carried over from the last period, the housing status with which the household enters period t, the house price, the wife's wage, the husband's income, the employment status, the household's nonlabor income, the individual match quality values and the Pareto weights $\tilde{\mu}_{g,t}$ (see below for more details on the latter).

Let $\mathbf{a}_t = (c_{M,t}, c_{F,t}, l_{F,t}, A_t, H_t, D_t)$ denote all the choices made by the household in any period t during the working phase. In addition, $\tilde{\mathbf{S}}_{M,t} = \left(A_{t-1}, H_{t-1}, p_t, y_{M,t}, e_{M,t}, y_t^{nl}\right)$ and $\tilde{\mathbf{S}}_{F,t} = \left(A_{t-1}, H_{t-1}, p_t, w_{F,t}, y_t^{nl}\right)$ are the relevant state variables for a divorced male and female at the start of period t.

The couple then solves the following problem:

[15] We do also refer the reader to Druedahl (2015), who studies portfolio allocation on different types of assets, but in the context of a unitary model and excluding divorce.

$$V_t^{mar}(\mathbf{S}_t) = \max_{\mathbf{a}_t} \left\{ \begin{array}{c} \tilde{\mu}_{M,t} u_M (c_{M,t}, H_t; \theta_{M,t}) + \tilde{\mu}_{F,t} u_F (c_{F,t}, l_{F,t}, H_t; \theta_{F,t}) \\ +\beta \mathbb{E}_t \left[\begin{array}{c} (1 - D_t) V_{t+1}^{mar}(\mathbf{S}_{t+1}) \\ +D_t \left(\tilde{\mu}_{M,t} V_{M,t+1}^{div} \left(\tilde{\mathbf{S}}_{M,t+1} \right) + \tilde{\mu}_{F,t} V_{F,t+1}^{div} \left(\tilde{\mathbf{S}}_{F,t+1} \right) \right) \end{array} \right] \end{array} \right\}$$

subject to the respective budget and borrowing constraints. This dynamic problem depends on the continuation value (discounted at rate β) and the utility in the given period. The Pareto weights determine the importance of both spouses in the decision process. A solution involves a mapping, say $\mathbf{a}^*(\mathbf{S}_t)$, from the state variables to decisions and determines the value of marriage for both spouses. Similar to Voena (2015), we can compute the continuation value through backwards recursion (see Online Appendix E for more details).

With respect to the continuation values, $V_{g,t}^{div}\left(\tilde{\mathbf{S}}_{g,t}\right)$ denotes spouse g's value of divorce. Households that choose to divorce need to liquidate debts and the net wealth is then distributed among ex-spouses via some division rule. To be more precise, before *White v. White* we assume a splitting rule that proxies for initial assets (at start of marriage), while we impose an equal split after the *White v. White* decision. In addition, even though we do not explicitly model remarriage in our structural model, we do include an additional calibrated term capturing the remarriage probability that is decreasing with age, so as to match realistic remarriage dynamics. The full expression for these value functions is given in Online Appendix D.

Given all this we obtain that spouse g's value of marriage is given by:

$$V_{g,t}^{mar}(\mathbf{S}_t) = u_g \left(c_{g,t}^*, l_{g,t}^*, H_t^*, \theta_{g,t} \right) + \beta \mathbb{E} \left[(1 - D_t) V_{t+1}^{mar}(\mathbf{S}_{t+1}) + D_t V_{g,t}^{div} \left(\tilde{\mathbf{S}}_{g,t+1} \right) \right], \quad (7)$$

where $c_{g,t}^*, l_{g,t}^*, H_t^*$ are part of $\mathbf{a}_t^*$, the optimal solution for the household in period t.

Participation constraints and Pareto weights. Due to limited commitment we assume that each spouse needs to specify the following participation constraint:

$$V_{g,t}^{mar}\left(\mathbf{S}_t\right) \geq V_{g,t}^{div}\left(\tilde{\mathbf{S}}_{g,t}\right). \tag{8}$$

That is, the value of staying inside marriage should be at least as large as the outside value of divorce, for each spouse g. For each period t there are three cases to consider. First, if the participation constraints for both spouses are satisfied, then nothing needs to change in terms of intra-household allocations. Next, if both spouses' participation constraints are violated, then there is no viable reallocation that makes both of them better off inside the marriage and they will decide to divorce. The interesting case is when only one spouse's participation constraint is not satisfied. In this case, the other spouse will have to give up resources in order to make the first one just indifferent between staying inside the marriage or divorcing. This can formally be captured using the Lagrange multiplier for spouse g's participation constraint, denoted by $\xi_{g,t}$. Following Marcet and Marimon (2019) one can then show that $\tilde{\mu}_{g,t+1} = \tilde{\mu}_{g,t}+\xi_{g,t}$. This makes that Pareto weight need not be constant over time (,i.e., limited commitment) and given our set-up this will change the resource allocation inside the household.

4 Identification and estimation

We start by listing the parameters that are preset. Subsequently we discuss the identification and estimation of the parameters outside and inside the model. We end by discussing the fit of our model.

4.1 Preset parameters

We follow the literature to preset the following parameters.

Discount factor, preference parameters and economies of scale. For the CRRA parameters we set $\sigma_M = \sigma_F = 2$ and for the discount factor we choose $\beta = 0.95$, which are typical values in the literature. Similar to Pizzinelli (2018), we set $\psi_F = 2.32$, implying a Frisch elasticity of labor supply of 0.5. The economies of scale parameter ρ is set to match the McClements scale, following Voena (2015), which leads to a value of 1.4023.

Initial assets. Given the general lack of good wealth information, we start the model for individuals aged 20 years old with no assets and starting out as renters. However, given that most of the (targeted) empirical moments apply to the age range of 30 and above, we allow individuals to accumulate some wealth before the effective start of the working phase.[16]

Prices of flats and renting cost. The parameter κ_F, reflecting the fraction of flat prices to house prices, is set to 0.6, which is the same value as in Attanasio et al. (2012). The rationale for this value stems from the average ratio of prices for homes with less than 5 rooms (incl. kitchen and bathrooms) to homes with more than 5 rooms. The rental price is set at 5 % of flat prices.

Credit market parameters. The rate of return on debt, R^- is set at 7 %, which approximately equals the average nominal interest rate on new mortgages in the UK in the early 2000s. The return on savings, R^+ is set to 3 %. The parameters determining the borrowing constraint (pertaining to the downpayment and debt-to-income limit) are chosen to reflect some specific UK institutional aspects. In particular, for λ^H we pick 0.9, which implies a downpayment of 10 %, which is reflecting the typical maximum value for the UK in the early 2000's (Pizzinelli, 2018). For the LTI-limit there is much more institutional variation, so we follow information in the *Guide to Mortgages* published by the Financial Services Authority in 2004.[17] This report

[16] In addition, households have access to nonlabor income from the start of the lifecycle.

[17] The FSA was a quasi-judicial body which accounted for regulating the financial industry in the UK between 2001 and 2013.

states that typically $\lambda^Y = 3$. We assume a (proportional) transaction cost of 7 % for selling and 2.5 % for buying, following Gruber and Martin (2004) and Yang (2009).[18]

Labor market parameters. The correlation between the productivity shocks among spouses, $\sigma_{\varepsilon_M,\varepsilon_F}$ is set to 0.25, which is the same as in Hyslop (2001) and Attanasio et al. (2018). To initialize the process, we picked a value of 0.2, which is the estimate of intra-household correlation of incomes in the year 2000 obtained by Lise and Seitz (2011). The unemployment probability π^u is set to 5.5 % and approximates the average unemployment rate in the UK in the early 2000s. The unemployment benefit is set to $b^u = 0.3$, the retirement income is given by a replacement rate of 50 % times the last income the individual earned during the working phase of the lifecycle. Age of retirement T_r is allowed to be asymmetric, where women retire from the age of 60, while men retire at 65.

Remarriage propensity parameter. The initial values of the remarriage propensity parameter, $\tilde{u}_g(t)$, are set at 0.9 for men and 0.5 for women. The difference for men and women is meant to capture the average lower remarriage rates for women. In each period t, we assume that divorcees receive an additional utility term representing an exogenous probability of remarriage. This gender-specific component declines linearly with age.

4.2 Parameters estimated and calibrated outside of the model

House prices. The variance of the house price shocks, σ_H^2 can be estimated using the second moment of growth rates in (net of trend) house prices. For this, we use data from the UK House Price Index (UKHPI) in the range 1968-2008. We deflated nominal house prices by the Retail Price Index (RPI, all items). The persistence

[18] The paper by Gruber and Martin uses data from the Consumer Expenditure Survey (CEX) on reallocation costs of tax and agency costs and find for median households costs of about 7 % for selling and 2.5 % for buying housing stock. Though these values represent the US context, they are on average close to the calibrated value of (symmetric) transaction costs equal to 5 % in Attanasio et al. (2018), who also use the BHPS to estimate their model.

coefficient, ρ_H is estimated to be equal to 0.87, while the standard deviation of house price shocks is given by 0.09.[19]

Wages and earnings. We used panel data on male earnings and female wages contained in the BHPS to estimate the deterministic lifecycle profiles for female wages $\alpha_F(t)$ and male earnings $\alpha_M(t)$, together with the variances for the permanent productivity shocks, $\sigma^2_{\varepsilon_M}$ and $\sigma^2_{\varepsilon_F}$.[20] To be more specific, we ran a regression of (log) earnings for men and (log) wages for women on age, age squared and other controls. We also control for selection into the labor market of women, using a Heckman two-step correction regression. The identification of the variances of the permanent shocks is based on the second moments of the growth rates of income and wages and is standard in the literature (Meghir and Pistaferri, 2004; Blundell et al., 2008).

Nonlabor income. To keep the size of the state space in our computations manageable, instead of adding an additional state variable, we used linear projections of (real) nonlabor income on a subset of the state space.[21]

Housing preference parameters singles. The parameters capturing the preferences for housing (homeownership and size) for singles are calibrated to help match moments pertaining to homeownership rates and median housing values for singles. More particularly, the parameter values are $\omega^{H,0}_{M,single} = 1.20, \omega^{H,1}_{M,single} = 1.24, \omega^{H,0}_{F,single} = 0.49$, and $\omega^{H,1}_{F,single} = 0.42$.

4.3 Parameters estimated within the model

We start by presenting how our parameters are indeed identified by our data. Subsequently we present our estimated parameters which are estimated via indirect infer-

[19] These values are quite similar to the estimates in Attanasio et al. (2012), who estimate the same house price process as we, but on a slightly different time period (1968-2000).

[20] See Online Appendix F for more details.

[21] The relevant state variables used are t (age) and $y_{M,t}$ for married individuals and single men, whilst we used t and $w_{F,t}$ for single women. Details of this estimation are provided in Online Appendix G.

ence.

4.3.1 Identification

Initial Pareto weight. The initial Pareto weights, $\tilde{\mu}_{M,0}$ and $\tilde{\mu}_{F,0}$, are estimated using exogenous variation in the post-divorce asset division in England due to the *White v. White* case. More specifically, the labor supply response of married women in Table 2 help to identify the (initial) bargaining power, similar to the identification strategy pursued in Voena (2015). Intuitively, for Pareto weights that are more favorable for men, the possibility for the household to reallocate resources is larger and therefore, we should expect a larger response in married women's reduction of working hours.[22]

Cost of working and preference for leisure. The cost of working for married women ($\delta_1^{mar}, \delta_2^{mar}$), and for single women ($\delta_1^{single}, \delta_2^{single}$) capture the hump-shaped lifecycle pattern of female employment rates, as illustrated in Figure 1. Intuitively, more concave profiles for the cost of working will force a hump shape pattern in the labor supply of married (resp. single) women. Average hours worked by married and single women across different age ranges and by marital status equally helps us with non-linear patterns in female labor supply on the intensive margin, but in addition add to the identification of the relative weight attached to leisure in married women's utility ($\omega^l_{F,mar}$) and for single women ($\omega^l_{F,single}$).

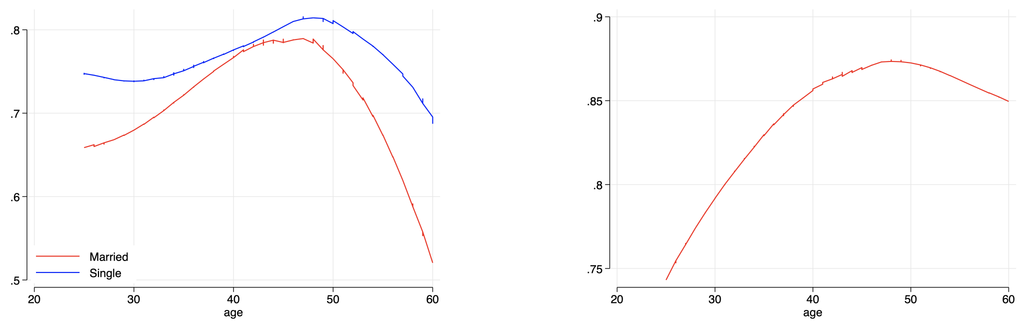
Housing preference parameters. The parameters $\omega^{H,0}_{g,mar}$ reflect the preferences for homeownership for married individuals, which directly influences the likelihood of households to buy a flat/house. Consequently, these parameters can be identified using homeownership rates for couples. In addition, the parameters $\omega^{H,1}_{g,mar}$ capture the preferences over housing size (that is, flat versus houses). These can be identified by the (median) housing values for both couples. Similar arguments hold for single

[22]This argument is also very similar to the analysis in Newman and Olivetti (2015), where they show that two-earner households can be more durable due to greater flexibility in terms of resources to reallocate such that both spouses are better off inside the marriage.

(wo-)men, for which we can then use the singles' counterparts for housing-related moments to identify $\omega_{g,single}^{H,0}$ and $\omega_{g,single}^{H,1}$.

Variance of match quality. The variances of match quality shocks $\sigma_{M,\theta}^2$ and $\sigma_{F,\theta}^2$ influences the likelihood of divorce. Higher volatility in match quality shocks imply a larger probability that at some point these shocks will trigger a divorce. Hence there is a theoretical link between these variance parameters and the average (crude) divorce rate.

Figure 1: *Employment and Homeownership rates over lifecycle.*



(a) *Female employment rate* (b) *Homeownership rate*

Notes: BHPS, respondents between 25 and 60 years old, sample range 1992-2005. Cross-sectional weights have been used.

4.3.2 Estimation of the auxiliary model using indirect inference

We use indirect inference (Gourieroux et al., 1993) to estimate the parameters that are not preset or estimated outside of the model. The main aspect of this method is the use of an auxiliary model, capturing important aspects of the data. This auxiliary model can be estimated both on observed data, as well as on simulated data from the dynamic structural model. Indirect inference then chooses the parameters in such a way to minimize the distance between these two alternative estimates of the auxiliary model. The auxiliary model to be matched contains the responses in labor supply and (crude) divorce rates to the *White v. White* case, female labor supply moments at different ages (on extensive and intensive margin) and by marital status and homeownership rates. These can be summarized in the vector $\boldsymbol{\phi}^{data}$. The estimates for the structural parameters, $\hat{\boldsymbol{\theta}}$, then solve the following:

$$\min_{\boldsymbol{\theta}} \left(\boldsymbol{\phi}^{sim}\left(\boldsymbol{\theta}\right) - \boldsymbol{\phi}^{data}\right)' \mathbf{W} \left(\boldsymbol{\phi}^{sim}\left(\boldsymbol{\theta}\right) - \boldsymbol{\phi}^{data}\right),$$

where $\boldsymbol{\phi}^{sim}\left(\boldsymbol{\theta}\right)$ denotes the simulated auxiliary model, and $\mathbf{W}$ is a symmetric, positive (semi-)definite weighting matrix.[23] Standard errors for $\hat{\boldsymbol{\theta}}$ are obtained from the asymptotic distribution presented in Gourieroux et al. (1993).

4.3.3 Parameter estimates and model fit

The parameter estimates are presented in Table 3. We obtain an initial Pareto weight of 0.74, which is a bit larger than the 0.7 found in Voena (2015). The linear coefficient capturing the (relative) preferences over leisure found in Pizzinelli (2018) is roughly in between our estimate for single and married women. Turning to our estimates of the monetary costs of female labor force participation, $\delta^{mar}(t)$ and $\delta^{single}(t)$, it is again useful to contrast our estimates to those found in Pizzinelli (2018). In particular, we find relatively similar (monetary) penalties for labor force participation of married women, whilst our estimates for single women is a bit larger. We also find preference reversals by marital status for both men and women. Indeed, while married women value homeownership (and housing size) more than single women, the opposite is estimated for men. Our structural approach is capable of matching the targeted moments in the data reassuringly well (see Table 16 in Online Appendix H for more details).

[23]We opted for the optimal weighting matrix, which coincides with the inverse of the variance-covariance matrix for the parameters of the auxiliary model, i.e., $W = \Sigma^{-1}_{\boldsymbol{\phi}^{data}}$.

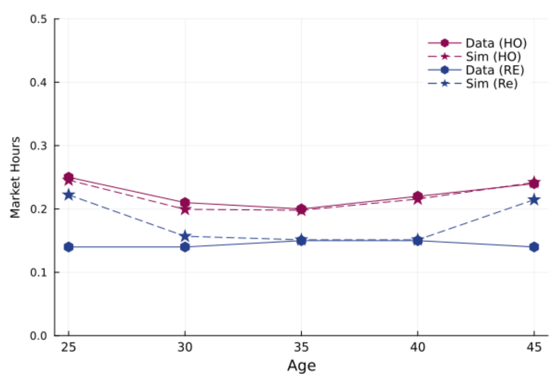
Table 3: *Parameter estimates.*

Parameter Description		*Estimate*	*Standard error*
Initial Pareto weight	$\tilde{\mu}_{M,0}$	0.74	0.207
Cost of work			
Cost of work married women (linear term)	δ_1^{mar}	0.026	3.3e-5
Cost of work married women (quadratic term)	δ_2^{mar}	0.0009	1.1e-6
Cost of work single women (linear term)	δ_1^{single}	0.076	1.48
Cost of work single women (quadratic term)	δ_2^{single}	0.0024	2.2e-7
Preference for leisure			
Preference for leisure married women	$\omega_{F,mar}^{l}$	0.26	0.001
Preference for leisure single women	$\omega_{F,single}^{l}$	0.1	0.0004
Preference for housing married individuals			
Utility homeownership men	$\omega_{M,mar}^{H,0}$	0.58	0.0016
Utility housing size men	$\omega_{M,mar}^{H,1}$	0.85	0.00073
Utility homeownership women	$\omega_{F,mar}^{H,0}$	1.1	0.001
Utility housing size women	$\omega_{F,mar}^{H,1}$	1.4	0.001
Match quality			
Variance match quality men	$\sigma_{M,\theta}^2$	0.169	0.000177
Variance match quality women	$\sigma_{F,\theta}^2$	0.15	0.00022

Notes: own calculations based on the same set of 10,000 simulations.

Out-of-sample fit. To further assess our model's performance, we compare several non-targeted moments from our simulations with their data counterparts. Figure 2 shows that our model can very well capture the fact that married women in homeowner households work more market hours on average than those in renting households. This finding is in line with earlier findings by Bottazzi et al. (2007) and Pizzinelli (2018).

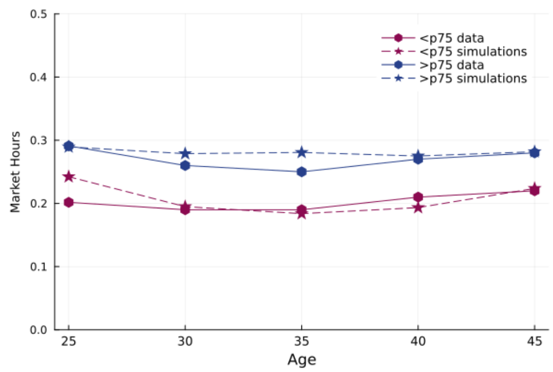
Figure 2: *Hours worked married women by housing tenure.*



Notes: own calculations based on the same set of 10,000 simulations.

Similarly, Figure 3 plots secondary earner hours against the primary loan-to-income (pLTI) ratio—computed by comparing households above and below the 75th percentile within each age group.[24] Our model replicates the finding that secondary earners in high-pLTI households work significantly more hours across all age groups. This pattern is consistent with the notion that higher pLTI ratios increase exposure to shocks to the primary earner's income, thereby requiring additional labor from the secondary earner as insurance.

Figure 3: *Hours worked married women by pLTI.*



Notes: own calculations based on the same set of 10,000 simulations.

Finally, Table 4 presents the average hours worked by married women, organized by 5-year age groups and quantiles of net housing wealth (defined as housing value minus outstanding debt).[25] Note that our simulated model matches quite well the expected negative gradient in the data between housing wealth and female labor supply, consistent with previous empirical studies (Henley, 2004; Milosch, 2014; Disney and Gathergood, 2018).

[24] The composition of those households below or above the 75th percentile is varying over the lifecycle. To take this into account, we computed the average hours worked in each age category broken down by the 75th percentile of the relevant leverage ratio within that same age group.

[25] We focus on the early part of the lifecycle between 30 and 54, given the absence of a realistic retirement dynamics in the model.

Table 4: *Hours worked married women by quantiles (net) wealth.*

Age	Q1 Data	Q1 Simulations	Q2 Data	Q2 Simulations	Q3 Data	Q3 Simulations	Q4 Data	Q4 Simulations
25 - 29	28	27	27	25	24	23	19	24
30 - 34	26	25	22	23	21	20	17	16
35 - 39	24	26	22	21	20	20	16	15
40 - 44	26	24	23	21	24	21	17	17
45 - 49	27	26	24	24	24	23	19	21

Notes: own calculations based on the same set of 10,000 simulations.

5 Responses to house price shocks

We now turn to the first main exercise to investigate the lifecycle dynamics of household responses to shocks in their housing wealth. More specifically, we will simulate a counterfactual model in which the prices are unexpectedly 10 % higher at a certain age and then revert back to their original trend values as given in equation (4), and then compare the resulting lifecycle path of the policy functions with their baseline outcomes. We will illustrate the aggregate effects on homeownership, female labour supply and divorce rates, as well as heterogeneous responses according to the change in housing status. Subsequently we use our model to further investigate the relationship between divorce rates, leverage ratios and match quality.

5.1 Aggregate effects

Figure 4 presents the average aggregate effects on homeownership, employment rate and divorce rates by age when the shock occurs. The sudden positive house price shock decreases housing demand, as access to the housing ladder has unexpectedly become steeper. This is reflected in a reduction of the homeownership rates at every age when the shock occurs. Given the dynamics in housing demand in the baseline model, it also intuitive that the house price shock has a stronger effect on homeownership rates at earlier ages than at later stages of the lifecycle. Due to the transitory nature of the shock, homeownership rates return to their baseline levels over time, though the shock does seem to have some persistence, in line with the persistent house price process.

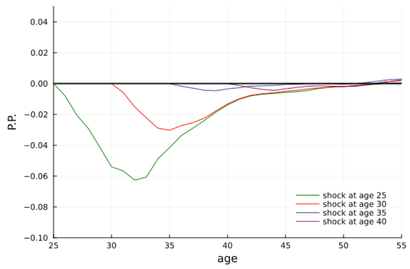
At the same time, the shock leads to a decline in married women's employment.[26] This reduction is driven by two mechanisms. For households that transition to renting (new renters), the easing of leverage-based borrowing constraints (notably the LTI limit) reduces the necessity for women to work. For those who remain homeowners, the appreciation in housing wealth increases the demand for leisure, thereby depressing labor supply—a result that aligns with the literature on unexpected wealth effects on household behavior (see, for instance, Cesarini et al. (2017);Golosov et al. (2023)). We also note that the responsiveness of employment rates is not monotonic; the largest drops in employment occur when the shock happens at age 35, while the responses for other ages are relatively similar. This pattern is largely driven by the baseline dynamics of married women's labor supply, particularly the U-shaped pattern where labor supply is at its lowest between ages 35 and 40—consistent with the (estimated) disutility of work for married women peaking around age 35.[27] This could imply that a drop in employment rates are then exacerbated by the additional housing wealth shocks.

Divorce rates go up in response to the positive house price shock, although the aggregate effect is ambiguous ex ante. For renters, the increased difficulty in accessing homeownership diminishes the expected lifetime wealth, thereby lowering the value of marriage. For homeowners, while housing wealth appreciation could, in principle, stabilize marriages, our results show that the negative impact of reduced expected housing wealth—through diminished marital surplus—dominates. In our model, this reduced expected housing wealth decreases the marital surplus, making spouses' participation constraints (8) more likely to bind. This finding contrasts with Rainer and Smith (2010), who document a protective wealth effect that lowers partnership dissolution. This is also corroborated by the fact that the responses of divorce rates closely track the responses of homeownership rates, which suggests that the loss of access to the housing ladder indeed has a destabilizing effect on households.

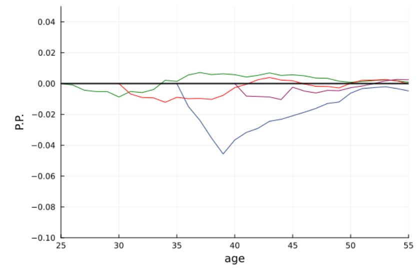
[26] In contrast to previous studies as in Low and Sánchez-Marcos (2024) that report a spurious positive association between house prices and female labor supply due to selection, our analysis shows that when selection and endogeneity are addressed, housing wealth shocks actually reduce employment.

[27] In reality, this pattern in the data is clearly driven by fertility episodes before that age range.

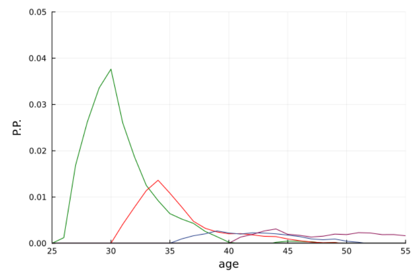
Figure 4: *Responses to an unexpected transitory increase in house prices.*



(a) *Homeownership rate*



(b) *Female employement rate*



(c) *Crude Divorce rate*

Notes: each plot presents the change (expressed percentage points) in respectively homeownership, female employment rate and crude divorce rate due to a 10% increase in house prices occurring at ages 25, 30, 35, and 40. All plots are produced with the same set of 10,000 simulations.

5.2 Heterogeneous responses according to change in housing status

While these aggregate results highlight the overall impact of unexpected house price increases, we may of course expect significant differences along the extensive margin of housing demand. Our structural model disentangles the underlying mechanisms—specifically, shifts in intra-household bargaining and wealth effects via borrowing constraints—by distinguishing between two groups: (i) "*Always homeowners*" (AH), which are those households that owned a property in both the baseline, as well as when the shock occurs, and (ii)"*New Renters*" (NR), which are households who owned a property in the baseline, whilst in the counterfactual scenario they are renting a property in response to the shock.[28] We examine lifecycle effects by analyzing

[28]The third group of households, consisting of those who were renting in both the baseline and in the counterfactual scenario, are less likely to access the housing ladder anyway. To streamline our discussion we therefore did not include them.

the same shock as before occurring at age 30 and age 40, focusing on the immediate responses in the period following the shock.

When the shock occurs at age 30, Table 5 shows that both groups experience a similar reduction in female labor supply, but the underlying drivers differ. In the NR-group, reduced housing demand lessens the reliance on costly female labor, resulting in an employment decline of approximately 6 percentage points. The observed increase in bargaining power for women in this group indicates that the decline in expected housing wealth and marital surplus renders their participation constraints more binding. Consequently, in couples that remain married, wives gain increased bargaining power. This also resulted in a substantial increase of the divorce rate, which indicates that no reallocation of resources could make both spouses better off. For the AH-group, the reduction in female labor supply is driven by a wealth effect. The effective lowering of the loan-to-value (LTV) limit allows these households to increase borrowing, which is reflected in reduced savings. This does not lead to a significant increase in binding participation constraints, which is reflected in mild changes in the female bargaining power and the divorce rate.

When the shock occurs later, at age 40, the qualitative results are largely consistent with those observed at age 30, with some notable differences. Specifically, there is no change in the divorce rate among the NR-group. However, marital instability still increases, as evidenced by a higher bargaining power for women. Similar to the age 30 shock, this increased bargaining power for women leads to a reduction in female labor supply. Additionally, households in the NR-group no longer need to rely on female labor to satisfy borrowing constraints, which further contributes to the significant drop in married women's employment rates.

It should be noted that the composition of the NR-group at age 40 differs from that at age 30. Indeed, Table 6 highlights the fact that the NR-group at age 40 is slightly lower in the income distribution compared to the AH-group as when the same comparison is made at age 30. This distinction is intuitive, as those households who now have to downsize and rent instead of buying a property are typically on the margin of being able to buy. At age 40, the NR-group predominantly consists of households that have only marginal access to the housing ladder, reflecting a lower ability to satisfy borrowing constraints.

Finally, we observe that the average spousal match quality among couples in the NR-group is higher than that in the AH-group—except for men when the shock occurs at age 40. This pattern suggests a selection effect in which only couples with sufficiently high match quality remain married when the material benefits of marriage (i.e., access to the housing ladder) are compromised.

Table 5: *Immediate responses to house price shocks by housing tenure.*

Housing status	Δ *divorce rate* (%)	Δ *bargaining power* (%)	Δ *employment rate* (p.p.)	Δ *savings* (%)
		Shock at age 30		
Always Homeowners	0.007	-0.088	-0.039	-0.056
New Renters	0.14	0.336	-0.062	-0.003
		Shock at age 40		
Always Homeowners	0.0065	-0.032	-0.034	-0.129
New Renters	0.0	0.265	-0.234	-0.195

Notes: own calculations based on the same set of 10,000 simulations.

Table 6: *Average match quality, earnings and wages by housing tenure and gender.*

Match quality men (θ_M)		*Match quality women* (θ_F)		*Earnings men* (y_M)		*Wages women* (w_F)	
AH	NR	AH	NR	AH	NR	AH	NR
			Shock at age 30				
-0.012	0.035	-0.005	0.065	0.842	0.637	0.922	0.794
			Shock at age 40				
-0.007	-0.128	-0.004	0.046	1.161	0.606	1.224	1.043

Notes: own calculations based on the same set of 10,000 simulations.

5.3 Divorce, leverage ratios and match quality shocks

So far, our analysis has focused on aggregated responses to exogenous, unexpected changes in housing wealth. We now broaden our investigation to address additional questions with important implications for marital stability and housing policy. In particular, we ask: do households with limited housing wealth or constrained credit face a higher risk of divorce, and through what mechanisms do these financial constraints affect intra-household bargaining and family outcomes?

Building on our reduced-form evidence in the Table 1 that household leverage ratios mediate the effect of house price shocks on divorce risk, we exploit simulated panel data from our dynamic structural model to further disentangle this relationship. In particular, we incorporate spousal match quality, as indicated by the differences observed in Table 6.

$$\Delta D_{h,t} = \beta_1 pLTI_{h,t-1} + \beta_2 LTV_{h,t-1} + \sum_{g\in\{M,F\}} \beta_{3,g} \mathbf{1}\left[\epsilon^{\theta}_{g,t} < 0\right] \times pLTI_{t-1}$$
$$+ \sum_{g\in\{M,F\}} \beta_{4,g} \mathbf{1}\left[\epsilon^{\theta}_{g,t} < 0\right] \times LTV_{t-1} + \beta_5 \mathbf{1}\left[H_{t-1} = 2\right] + \beta_6 \mathbf{1}\left[H_{t-1} = 3\right] + \epsilon_{h,t},$$

where $\Delta D_{h,t} = D^{\text{countfac}} - D^{\text{baseline}}$ equals one if a shock-induced divorce occurs and zero otherwise. Our key regressors—the pre-shock debt-to-primary income ratio ($pLTI$) and the debt-to-value ratio (LTV)- are lagged values to study the effect of the beginning of the period indebtedness and the loan-to-value ratio on the divorce decisions. These regressors are interacted with dummies for a negative match quality shock, i.e., $\mathbf{1}\left[\epsilon^{\theta}_{g,t} < 0\right]$, with $\epsilon^{\theta}_{g,t} = \Delta\theta_{g,t}$. Finally, we also control for housing tenure and estimate this regression for shocks at ages 30 and 40. The results are presented in Table 7.

For both ages, our findings confirm the reduced-form evidence (Table 1). Households with high $pLTI$ ratios are more likely to divorce in response to a house price shock—consistent with tighter borrowing constraints limiting the ability to upscale housing and reducing marital surplus. In contrast, higher LTV ratios, which relax these constraints, are associated with lower divorce risk. These effects are even stronger when the shock occurs at age 40, suggesting that at later life stages, households are particularly sensitive to intensive margin adjustments on the housing ladder.

Moreover, our analysis reveals that spousal match quality plays a moderating role. Adverse match quality shocks reduce the nonmaterial benefits of marriage by tightening spouses' participation constraints. Specifically, in high $pLTI$ households, negative match quality shocks for wives amplify divorce risk, whereas in high LTV households the effect is mitigated by greater borrowing capacity. For husbands, a negative match quality shock can increase the wife's bargaining power—due to lower relative earnings—further destabilizing the marriage. Notably, these dynamics are more pronounced when the shock occurs at age 40.

To corroborate these findings, we estimate a similar regression using the change in the female Pareto weight,

$$\Delta\tilde{\mu}_{F,t} = \tilde{\mu}_F^{\text{countfac}} - \tilde{\mu}_F^{\text{baseline}},$$

as the dependent variable (Table 8). The results indicate that in high $pLTI$ households, increased divorce risk is accompanied by a significant revision of bargaining power in favor of the wife, whereas in high LTV households no significant change in the wife's bargaining power is observed, consistent with the lower divorce likelihood.

Table 7: *Responses to transitory house prices: divorce and leverage ratios.*

	ΔD (*Shock age 30*)	ΔD(*Shock age 40*)
$pLTI_{t-1}$	0.0013***	0.0024***
	(0.0002)	(0.0002)
LTV_{t-1}	-0.006***	-0.0130***
	(0.001)	(0.0009)
$\mathbf{1}\left[\epsilon^{\theta}_{M,t} < 0\right] \times pLTI_{t-1}$	-0.0003	-0.0010***
	(0.0002)	(0.0002)
$\mathbf{1}\left[\epsilon^{\theta}_{F,t} < 0\right] \times pLTI_{t-1}$	0.0009***	0.00151***
	(0.0002)	(0.0002)
$\mathbf{1}\left[\epsilon^{\theta}_{M,t} < 0\right] \times LTV_{t-1}$	0.0013	0.0036***
	(0.001)	(0.001)
$\mathbf{1}\left[\epsilon^{\theta}_{F,t} < 0\right] \times LTV_{t-1}$	-0.0020*	-0.0043***
	(0.001)	(0.001)
$\mathbf{1}\left[H_{t-1} = 2\right]$	0.0020***	0.0010***
	(0.00029)	(0.00027)
$\mathbf{1}\left[H_{t-1} = 3\right]$	-0.0001	-0.0012***
	(0.0003)	(0.0003)
Observations	300,621	306,928
R-squared	0.0009	0.0037
Demographic controls	✓	✓

Notes: regression on same set of 10,000 simulations. The set of demographic controls consists of the age of the household and primary earnings.

Table 8: *Responses to transitory house prices: bargaining power and leverage ratios.*

	$\Delta\tilde{\mu}_F$ *(Shock age 30)*	$\Delta\tilde{\mu}_F$*(Shock age 40)*
$pLTI_{t-1}$	0.0031***	0.0022***
	(0.0003)	(0.0003)
LTV_{t-1}	-0.0026*	0.0004
	(0.001)	(0.001)
$\mathbf{1}\left[\epsilon^{\theta}_{M,t} < 0\right] \times pLTI_{t-1}$	-0.0005*	-0.0005*
	(0.0003)	(0.0003)
$\mathbf{1}\left[\epsilon^{\theta}_{F,t} < 0\right] \times pLTI_{t-1}$	-0.0001	0.0001
	(0.0003)	(0.0003)
$\mathbf{1}\left[\epsilon^{\theta}_{M,t} < 0\right] \times LTV_{t-1}$	0.0021	0.0013
	(0.001)	(0.001)
$\mathbf{1}\left[\epsilon^{\theta}_{F,t} < 0\right] \times LTV_{t-1}$	0.0005	0.00001
	(0.001)	(0.001)
$\mathbf{1}\left[H_{t-1} = 2\right]$	-0.006***	-0.001***
	(0.0004)	(0.0003)
$\mathbf{1}\left[H_{t-1} = 3\right]$	-0.0143***	-0.007***
	(0.0004)	(0.0004)
Observations	300,621	306,928
R-squared	0.005	0.006
Demographic controls	✓	✓

Notes: regression on same set of 10,000 simulations. The set of demographic controls consists of the age of the household and primary earnings.

6 Responses to credit market tightening

Given the reported results in the last section on the importance of leverage ratios in transmitting shocks to marital instability, we conduct a policy counterfactual by tightening the income-related borrowing constraint. Specifically, we reduce λ^Y from 3 to 2.5 in equation (6) and study the long-run impact on households of such a tightening in the credit market.[29]

As a consequence, households are forced to adjust their behavior. First, with a

[29]For the period under consideration 2.5 is on the lower yet realistic end of the LTI-limit. Indeed, as the Financial Services Authority guide to mortgages in 2004 states: *"Typically, the maximum mortgage a lender offers is three times the main earner's income plus one times any second earner's income, or two-and-a-half times your joint income."* This number is also quoted in Pizzinelli (2018).

lower borrowing capacity, they may rely more on the secondary earner's labor supply in order to afford property purchases. Second, a reduction in borrowing capacity curbs expected housing wealth accumulation and thus diminishes the marital surplus. In our model, this reduction in surplus makes spouses' participation constraints (8) more likely to bind, thereby increasing the risk of divorce.

An important strength of our model is that it allows us to assess the roles of divorce and limited commitment channels in transmitting this credit market tightening to household outcomes and overall welfare. To illustrate this, we consider three scenarios. The first is our baseline model, where spouses interact under *Limited Commitment (LC)* and can adjust bargaining weights when a participation constraint binds. The second scenario, *Full Commitment (FC)*, fixes spouses' bargaining power from the start of the marriage and prohibits divorce, effectively yielding a unitary household model as in Chiappori and Mazzocco (2017). Finally, in the *Full Commitment with Divorce (FC & D)* scenario, bargaining power remains fixed as in FC, but divorce is allowed. Notably, while our results are broadly consistent with Pizzinelli (2018), they also indicate that intra-household adjustments can significantly modify the macroeconomic impact of leverage constraints.

6.1 Aggregate effects

Table 9 reports the long-run responses of key outcomes—homeownership rates, divorce rates, bargaining power, and employment—to a tightening of the income-related borrowing constraint.[30]

Across all scenarios, homeownership rates decline by roughly 10 percentage points. This decline lowers the marital surplus, which, in turn, triggers intra-household renegotiations when one or more participation constraints bind. In fact, Table 9 shows that under the baseline *Limited Commitment* scenario, women's participation constraints are more likely to bind, consistent with the higher estimated marginal utility for housing among married women (see Table 3). As a result, women's Pareto weights increase, which leads to greater leisure and a dampened labor supply response to the

[30]By "long run" we indicate the overall average effect from age 30 until retirement.

tightening of the borrowing limit. The reduced labor supply response further reinforces the housing demand channel, yielding the largest drop in homeownership under LC.

By contrast, in the *Full Commitment* scenario—where bargaining weights are fixed and divorce is not permitted—the household cannot adjust its resource allocation in response to the tighter borrowing constraint. Consequently, the labor supply channel is more responsive, partly offsetting the decline in homeownership (which drops by about 9.5 percentage points). Finally, in the *Full Commitment with Divorce* scenario, although bargaining weights remain fixed, divorces are allowed. Here, selection effects are important: households that might have compensated through resource reallocation under LC instead divorce, and these households tend to have lower income. As a result, the drop in homeownership is even smaller (around 8.9 percentage points), and the employment response is somewhat weaker compared to the FC scenario. Notably, while our results are broadly consistent with Pizzinelli (2018), they also indicate that intra-household adjustments can significantly modify the macroeconomic impact of leverage constraints.

Table 9: *LTI tightening: long run outcomes*

	Δ *homeownership* (p.p)	Δ *divorce rate* (%)	Δ *bargaining power* (%)	Δ *employment rate* (p.p.)
LC	-0.13	0.015	0.064	0.001
FC	-0.095	-	-	0.03
FC&D	-0.089	0.023	-	0.022

Notes: own calculations based on the same set of 10,000 simulations.

6.2 Heterogeneous responses according to change in housing status

As in Section 5, we again report the heterogeneous responses along the extensive margin of housing demand. Recall that we classify households into two groups at age 30: *Always Homeowners* (AH) and *New Renters* (NR). Recognizing that intra-household bargaining is a critical channel through which households adjust to tighter credit conditions, we further explore how these bargaining shifts are reflected in changes in private consumption. Table 10 presents our findings on consumption adjustments alongside female employment across tenure groups.

Under Limited Commitment, the labor supply response is notably dampened for both groups compared to the FC and FC & D scenarios. This suggests that increased bargaining power among women—partly due to their higher disutility of labor—leads to a lower overall employment response. In fact, women in the NR group experience an average increase in private consumption of about 4%, while their husbands see a drop of roughly 11%. Such resource reallocation is consistent with the enhanced female bargaining position observed under LC.

We also note that for the other two models excluding rebargaining leads to a counterfactual in which households in both groups tend to mainly respond by increasing their savings to adjust for the tighter credit conditions. This is in particular true for the NR-group.

Table 10: *Breakdown of long run responses*

	Δ *employment rate* (p.p.)	Δ *female consumption* (%)	Δ *male consumption* (%)
		LC	
Always Homeowners	0.01	0.041	0.015
New Renters	0.016	0.044	-0.112
		FC	
Always Homeowners	0.058	-0.006	-0.006
New Renters	0.035	-0.082	-0.082
		FC & D	
Always Homeowners	0.039	0.004	0.004
New Renters	0.039	-0.064	-0.064

Notes: own calculations based on the same set of 10,000 simulations.

6.3 Welfare costs of LTI tightening

As a final exercise, we evaluate the individual welfare effects of the LTI tightening using a consumption-equivalent variation measure, which calculates the proportional change in private consumption needed to make individuals indifferent between the baseline and counterfactual economies.[31] Given the tighter LTI limit—restricting credit access—we expect negative welfare effects. This is particularly true for the "New Renters". For "Always Homeowners", welfare losses arise from higher down-

[31] See Online Appendix, Section I for details on the equivalent variation calculations.

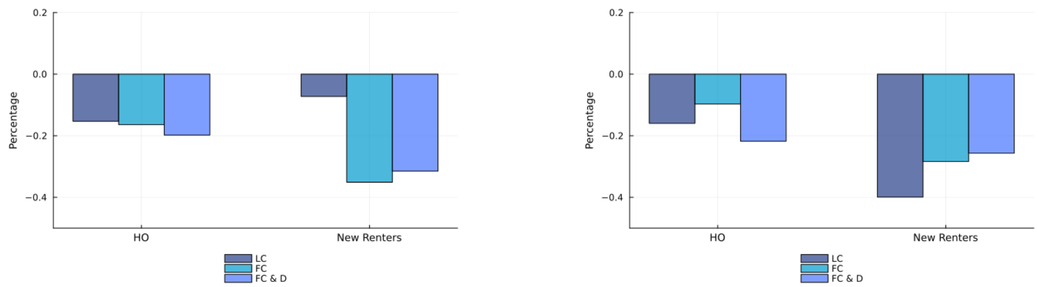
payment requirements and a forced downsizing of housing demand (i.e., households that could afford a larger house in the baseline can only purchase a smaller property under the new limit).

Our model enables us to quantify the contributions of intra-household bargaining and divorce channels to these welfare effects. Figure 5 presents the welfare costs separately for married men and women across various specifications. Under the baseline *Limited Commitment* scenario, overall welfare losses are highest for New Renters. Notably, married women in this group experience almost no welfare loss (approximately 7.241% change), while married men incur the largest welfare cost (approximately 39.9% change). This gender difference is consistent with our finding that women's participation constraints bind more frequently under LC, allowing them to enhance their bargaining power—either through renegotiation or, if renegotiation is not feasible, through divorce—which in turn mitigates their welfare losses.

Since men work full-time, our welfare measure focuses on proportional changes in private consumption to have a fair comparison. We refer to Online Appendix J for a Hicksian equivalent variation that incorporates adjustments in *both* consumption and leisure.[32] The qualitative results remain similar: the full model predicts substantially lower welfare costs for married women compared to scenarios where the intra-household bargaining and divorce channels are disabled. Once more this demonstrate the importance of limited commitment.

[32] In addition, we also calculated the welfare costs for married women when we fix their labor supply at the baseline levels, thereby completely excluding any compensation through that channel. These results are also presented in Online Appendix J.

Figure 5: *Welfare costs of LTI-tightening.*



(a) *Married women* (b) *Married men*

Notes: own calculations based on the same set of 10,000 simulations.

7 Conclusion

We analyzed how house price shocks affect marital stability and household decisions, in particular labor supply. To do this, we first showed reduced-form evidence that positive house price shocks, in combination with high pLTI levels (capturing indebtedness of the household relative to the primary earner's income) increases divorce risk, whilst households with high LTV levels experience null or even slightly decreased effects on their probability to divorce.

To understand the mechanisms behind these results, we then presented and estimated a dynamic model of the household. The main mechanism behind the reduced-form results are the fact that the decrease in expected housing wealth reduces marital surplus, which makes it more likely that the participation constraint of at least one spouse, notably women, becomes more binding. This effect is particularly the case in households with high pLTI ratios, as in these households the borrowing limit is even tighter, meaning that the increased servicing cost of debt falls on the wife's labor supply. In contrast, high LTV households experience a loosening of the borrowing constraint and therefore an increase in the expected marital surplus.

We then conducted a counterfactual policy analysis of a tightening in the debt-to-income limit in the credit market. We have shown how intra-household bargaining and divorce are important channels to understand the responsiveness of homeownership and married women's employment rates at the aggregate level, as well as to conduct welfare analysis of such credit market policies.

We believe that our rich structural model of how house prices shock interact with household decisions is the beginning of an interesting research agenda that focuses on different research questions and offers scope for further extensions or our approach. One such extension is the inclusion of raising children and fertility decisions, which is tied to both housing demand and labor supply. Two alternative examples are giving a proper treatment of the retirement stage in order to better grasp the insurance aspects and wealth effects of marriage related to this phase of life. Or, while divorce was endogenous in our approach, we did not model the marriage market (in particular, choice of one's partner). While all this received ample attention in the literature, it has not been done in a rich context as ours. As such this offers scope for further improving our understanding of how changes in the value of marriage impact household behavior.

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Appendix (for online publication)

A Details on house price shock measure

To construct the house price shock measure, $HPshock_{l,t}$ at local authority (LA) l in year t, we first estimate an AR(2) regression of (log) average prices of semi-detached houses, $\ln(AvePrice_{l,t})$, controlling for a set of LA-level fixed effects. The results are as follows:

Table 11: *AR(2) estimates for LA -level prices*

	$\ln(AvePrice_{l,t})$
$\ln(AvePrice_{l,t-1})$	1.42***
	(0.0089)
$\ln(AvePrice_{l,t-2})$	-0.61***
	(0.0090)
Observations	8,263
R-squared	0.99
LA FE	✓

Notes: estimation on Halifax house price data, sample range 1995-2017.
House prices are defined as average of prices of semi-detached houses.

Reassuringly, these results are very similar to the ones obtained in Disney et al. (2010). From this, we then accumulated the residuals of this regression over the past 3 years, which by construction yields a relatively symmetric distribution with a mean very close to zero.

B Further institutional background on *White v. White* case

England and Wales have, with some generalization, an equitable division system, in the sense that courts have a large amount of discretion in deciding splits of property and wealth among spouses upon divorce. This is an important difference with other countries, e.g. France or Italy, where people can choose between a particular default property regime and an alternative one. Furthermore, as noted by Smith (2003) and Piazzalunga (2017), pre-marital contracts regarding division of property are also quite uncommon in the United Kingdom, given that such contracts are not legally binding. Both of these aspects in divorce law are quite useful for our own analysis in this paper, since we do not have to be specifically concerned about selection effects of individuals into particular property division regimes.

Instead, ex-post agreements between spouses are encouraged. When spouses agree on such a division, the proceedings for divorce are simplified in the sense that the court just needs to issue a cheaper 'consent order'. In contrast, when the spouses can't agree on such a settlement, the court needs to issue a financial order, which takes more time and is more expensive.[33] From a practical perspective, the division rule as applied before the *White v. White* case was such that, when wealth (matrimonial assets) exceeded the financial needs of the household members, the remaining proceedings were distributed on a 'needs-based' system, that is, taking into account the specific financial needs of the former spouses and their standard of living they were used to.[34]

As noted in Piazzalunga (2017), who cites the *Ferguson v. Ferguson* case from 1994, the court in that case described such an equitable regime as "more fair" than a titular-based system. However, in most cases the courts didn't grant much larger shares to the wife, except for special cases (e.g. in the case where both partners were also business partners), and the wife didn't receive a share larger than 50 % in combination with much smaller shares of previously joint assets (Smith, 2003; Piazzalunga, 2017). A particular example of the latter is the so-called *Dart v. Dart*

[33]Though the option of agreeing ex post on a division rule is important, from an economic perspective the relevant factor is still the (credible) outside option of what each spouse could obtain through a financial settlement enforced by the courts. If the latter employ a more egalitarian division of property, then this would benefit the economically weaker spouse at the point of bargaining.

[34]In practice, courts made (and in some cases still make) use of the so-called 'Duxbury Tables' to calculate the 'reasonable needs' of each spouse. In particular, it is a lump sum amount that is calculated based on the assumption that the economically weaker spouse spends a share of capital and interest received in such a way that when (s)he dies, there is no capital left.

case in 1996.[35]

C Validity of our Difference-in-Difference approach

Several of the exercises in this appendix are similar to the approach taken by Piazzalunga (2017). We refer to that paper for a more extensive analysis of the empirical results (including a discussion on simultaneous policy changes occurring with the White v. White case).

C.1 Summary statistics

Table 12: *Summary statistics of the regression sample.*

Variable	Mean	St. deviation
Age	43.13	9.67
Higher educated	0.30	0.46
Wage	9.79	8.53
Non-labor income	4,506.76	6,513.43

Notes: summary statistics for our main estimation sample of married women in the age range of 25-60 years old during the period 1992-2005. We have a total of 27,589 observations. Non-labor income and wage levels are deflated by the CPI with 2014 as reference year. Higher educated refers to having at least obtained A-levels. Cross-sectional weights have been used.

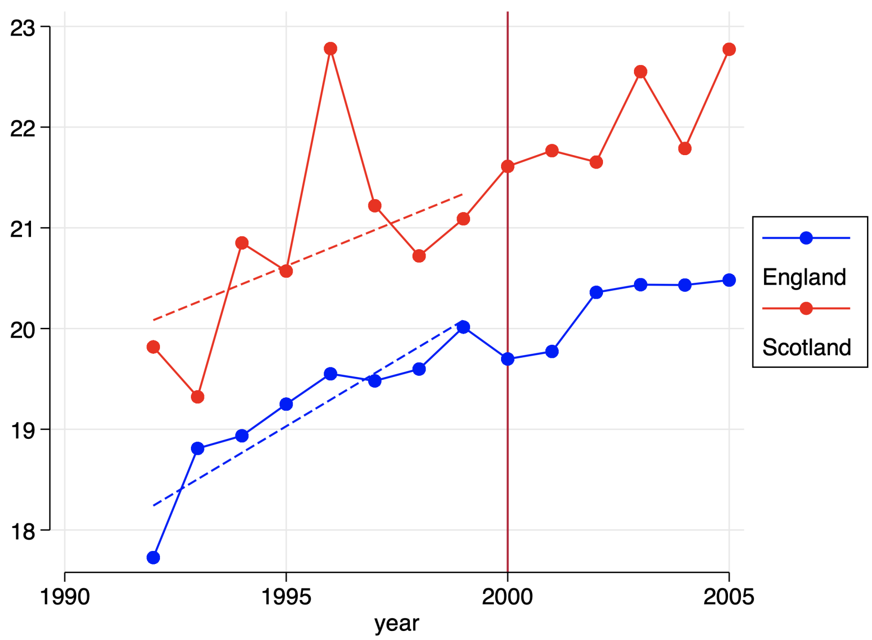
C.2 Standard tests

Parallel trends. We formally test for the parallel trends assumption on our main outcome variables (in particular married women's hours worked and employment status), by regressing the outcome variables on a full set of interactions between a dummy indicating whether the respondent is living in Scotland and the year dummies.

[35] Mr. and Mrs. Dart moved to England from the US (Kentucky), but were living in England when the wife filed for divorce. Mrs. Dart tried to get the case settled in the US, but eventually the case was decided in English courts. The stakes were quite large, given that Mr. Dart had a large fortune estimated to 400 million GBP. Mrs. Dart sought to get a settlement at around 100 million GBP. However, she lost both at the High Court and the Court of Appeal and eventually only got awarded 8.5 million GBP and had to pay the legal costs of her husband.

We then test common trends by statistically testing the null hypothesis that the interaction effects are null for the years before 2000 against the alternative hypothesis they are not null. This is similar to the approach taken by Ohinata and Picchio (2020). The p-value for the test with *Hours* equals 0.15, and for *Employment* it is 0.86. Hence, we cannot reject the null hypothesis and therefore we cannot reject the common trends assumption. For completeness we also show the pre-*White v. White* trends for labor supply here:

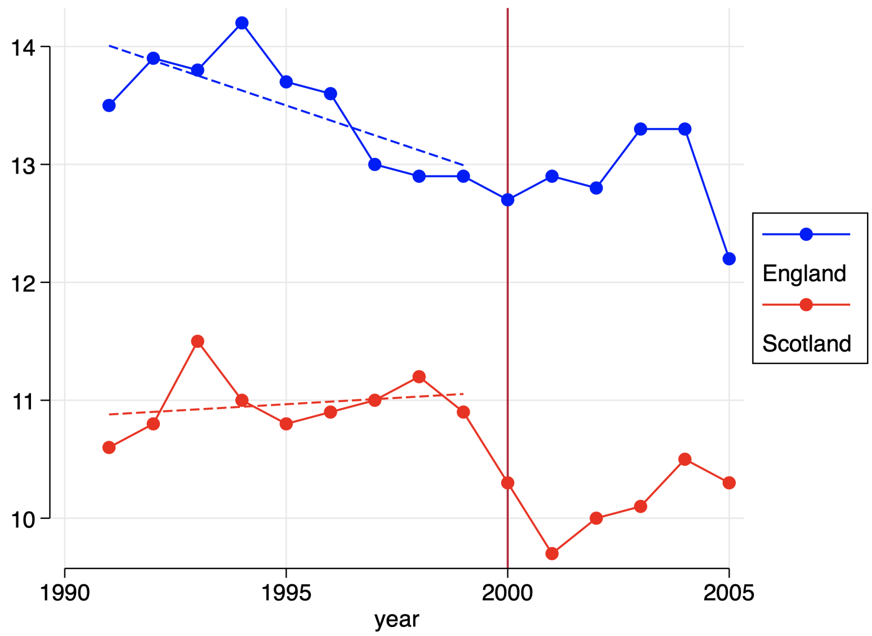
Figure 6: *Hours worked married women.*



Notes: BHPS sample, married women in the age range of 25-60 years old during the period 1992-2005.

Due to the limited number of observations, we cannot directly use the same statistical test of parallel trends for the crude divorce rates, but we present the pre-*White v. White* trends here below:

Figure 7: *Crude divorce rates.*



Notes: Office for National Statistics , 'Vital statistics: Population and Health Reference tables', sample range 1990-2005.

Placebo tests. To verify for the validity of the DiD estimates, we conduct a placebo test by checking for any labor supply effects of *White v. White* on women who are either cohabiting, never married or single (never married or divorcees) throughout our years of observations. For all these groups, there should not be any discernible labor supply responses, given that these are not treated by the case. Table 13 presents these results and confirm that there are no significant labor supply responses at the intensive margin. We also found again no effect at the extensive margin.

Table 13: *White v. White and labor supply of cohabiting and single women.*

	$Hours_{i,c,t}$	$Hours_{i,c,t}$	$Hours_{i,c,t}$
$Post \times Treated_{c,t}$	3.32	-0.93	-0.63
	(2.48)	(1.89)	(1.37)
Observations	4,942	4,479	9,486
R-squared	0.24	0.36	0.35
Sample	Cohabiting	Never married	Never married & divorcees
Demographic (incl. spouse) controls	✓	✓	✓
Year	✓	✓	✓
Region	✓	✓	✓
Local female unemployment rate	✓	✓	✓

Notes: estimation on BHPS sample, never married, divorcees and cohabiting women in the age range of 25-60 years old during the period 1992-2005. Standard errors clustered at the individual level, *** p<0.01, ** p<0.05, * p<0.1. Demographic controls include the age and age squared of husband and wife, the education level of husband and wife both defined as their highest qualification received, number of young children ($\leq$ 15 age old) and the partner's non-labor income in case of cohabitation. Non-labor income is deflated by the CPI with 2014 as reference year. Cross-sectional weights have been used.

C.3 Extensions and robustness

Male labor supply & Panel estimation. We can also estimate the effect of the *White v. White* case for married men. From a theoretical perspective, given that the case induced a higher bargaining power for women, together with the fact that a majority of married men work full time, we do not expect any significant changes in male labor supply. This is indeed confirmed in Table 14. We also re-estimated our main DiD specification as represented in (1) by exploiting the panel dimension in the data, which allows us to include individual fixed effects to control for unobserved heterogeneity that does not change over time. To do this, we only preserve those women in the sample which we observe at least once before the year 2000 and once after. The results are also presented in Table 14. The qualitative results are similar to the DiD results in Table 2. The coefficient of the effect of *White v. White* is estimated with more noise and the point estimate reduces in size (to a reduction of about 1.4 hours worked per week). We also conducted the same regressions for employment status of married men and employment status of married women incl. individual FEs, but did not find any significant effects.

Table 14: *White v. White: extensions and robustness.*

	$Hours_{i,c,t}$	$Hours_{i,c,t}$
$Post \times Treated_{c,t}$	-0.57	-1.41*
	(1.16)	(0.899)
Observations	23,330	24,232
R-squared	0.14	0.098
Sample	Married men	Married women
Individual FE		✓
Demographic (incl. spouse) controls	✓	✓
Year	✓	✓
Region	✓	✓
local male unemployment rate	✓	✓

Notes: estimation on BHPS sample, respectively married men and women in the age range of 25-60 years old during the period 1992-2005. Standard errors clustered at the individual level, *** p<0.01, ** p<0.05, * p<0.1. Demographic controls include the age and age squared of husband and wife, the education level of husband and wife both defined as their highest qualification received, number of young children ($\leq$ 15 age old) and the household's non-labor income. Non-labor income is deflated by the CPI with 2014 as reference year. Cross-sectional weights have been used.

Household work. The more egalitarian distribution of assets upon divorce following *White v. White* has had a negative effect on married women's labor supply. Given that our structural model does not include domestic work, we need to know whether there is any evidence of substitution of this freed up time towards domestic work. The BHPS asks respondents how much time they spend (on average) per week on domestic chores, which we can use as a proxy for household work. We then run a regression of household work on the same set of controls as in our main DiD specification as represented in (1). Table 15 presents the results for both the OLS and fixed effects regression.

Table 15: *White v. White and household work.*

	Household work$_{i,c,t}$ OLS	*Household work*$_{i,c,t}$ Fixed effects
$Post \times Treated_{c,t}$	1.284**	0.004
	(0.555)	(0.515)
Observations	26,853	24,415
R-squared	0.129	0.056
Individual FE		✓
Demographic (incl. spouse) controls	✓	✓
Year	✓	✓
Region	✓	✓
local female unemployment rate	✓	✓

Notes: estimation on BHPS sample, married women in the age range of 25-60 years old during the period 1992-2005. Standard errors clustered at the individual level, *** $p<0.01$, ** $p<0.05$, * $p<0.1$. Demographic controls include the age and age squared of husband and wife, the education level of husband and wife both defined as their highest qualification received, number of young children ($\leq$ 15 age old) and the household's non-labor income. Non-labor income is deflated by the CPI with 2014 as reference year. Cross-sectional weights have been used.

Though the pooled OLS estimate in column 1 suggests am increase in domestic work (hence suggestive for a substitution between market and domestic work time), this effect vanishes (both in size and statistically) after controlling for individual fixed effects, i.e., unobserved heterogeneity at the individual level. All in all this suggests that not including domestic work should not be too influential.

D Further details on the structural model

D.1 Splitting rule

The model captures the idea that pre-*White v. White* the distribution of assets was more unequal. To reflect this, we use the following split rule for assets upon divorce:

$$A_{M,t} = \tfrac{y_{M,t}}{y_{M,t}+w_{F,t}\times\bar{n}_F} A_t \qquad \text{and} \tag{9}$$

$$A_{F,t} = \left(1 - \tfrac{y_{M,t}}{y_{M,t}+w_{F,t}\times\bar{n}_F}\right) A_t, \tag{10}$$

before the *White v. White* case and an equal split afterwards. Here $\bar{n}_F$ denotes an average level of women's labor supply and we thus treat relative (potential) earnings for spouses as a proxy for a static bargaining rule upon divorce. Although more elaborated than standard in the literature, this is of course a simplification, mostly due to a lack of information and good data on either individualized wealth or any voluntary ex-post settlements between (former) spouses.

D.2 Optimization problems for singles

Preferences. Single men and women are very similar to their resp. married counterparts:

$$u_g^{single}\left(c_{g,t}, l_{g,t}, H_t\right) = \frac{\left(c_{g,t}\right)^{1-\sigma_g}}{1-\sigma_g} + \omega_{g,single}^{l}\frac{\left(l_{g,t}\right)^{1-\psi_g}}{1-\psi_g} + \omega_{g,single}^{H}\mathbf{1}\left[H_t > 1\right].$$

Note that singles do not experience match quality shocks and that we allow for the preferences for leisure to be different between single and married women (that is, we allow $\omega_{F,single}^{l} \neq \omega_{F,mar}^{l}$).

Budget and borrowing constraints. Single women face the following budget constraint:

$$\begin{aligned} & A_t + c_{F,t} + \tilde{p}_t(H_t)H_t + \Omega\left(H_t, H_{t-1}\right) \\ = & \left(1+R\right)A_{t-1} + w_{F,t}n_{F,t} + y^{nl}_{F,t} - \delta^{single}\left(t\right)\mathbf{1}\left[n_{F,t} > 0\right] \\ & + \tilde{p}_t\left(H_{t-1}\right)H_{t-1}\mathbf{1}[H_{t-1} > 1]. \end{aligned} \tag{11}$$

Note that obviously singles no longer have economies of scale in private consumption. We also allow for different monetary cost of participating on the labor market, compared to their married counterparts (that is, we allow $\delta^{single}(t) \neq \delta^{mar}(t)$). This is motivated by differential employment rates of women by marital status.

Similarly, single men their budget constraint is given by:

$$\begin{aligned} & A_t + c_{M,t} + \tilde{p}_t(H_t)H_t + \Omega\left(H_t, H_{t-1}\right) \\ = & \left(1+R\right)A_{t-1} + Inc_{M,t} + \tilde{p}_t\left(H_{t-1}\right)H_{t-1}\mathbf{1}[H_{t-1} > 1]. \end{aligned} \tag{12}$$

Finally, the borrowing constraint for single women is given by

$$Debt_t \leq \max\{Debt_{t-1}, \min\{LTV_t, LTI_t\}\} \tag{13}$$

$$LTV_t = \lambda^H \tilde{p}_t\left(H_t\right),$$

$$LTI_t = \lambda^Y n_{F,t} \times w_{F,t},$$

and for single men by

$$Debt_t \leq \max\{Debt_{t-1}, \min\{LTV_t, LTI_t\}\} \tag{14}$$

$$LTV_t = \lambda^H \tilde{p}_t\left(H_t\right),$$

$$LTI_t = \lambda^Y Inc_{M,t}.$$

Optimization problems. Let $\mathbf{a}^{single}_{F,t} = (c_{F,t}, l_{F,t}, A_t, H_t)$ and $\mathbf{a}^{single}_{M,t} = (c_{M,t}, A_t, H_t)$ denote the singles' choice variables during the working phase $t \leq T_r$. We then obtain the following optimization problem for individual g

$$V_{g,t}^{single}\left(\tilde{\mathbf{S}}_{g,t}\right)=\max_{\mathbf{a}_{g,t}^{single}}u_{g}^{single}\left(c_{g,t},l_{g,t},H_{t}\right)+\beta\mathbb{E}\left[V_{g,t+1}^{single}\left(\tilde{\mathbf{S}}_{g,t+1}\right)\right],$$

subject to (11) and (13) for women or (12) and (14) for men.

D.3 Value of divorce

The value function for divorcee g is given by:

$$V_{g,t}^{div}\left(\tilde{\mathbf{S}}_{g,t}\right)=\max_{\mathbf{a}_{g,t}^{div}}u_{g}\left(c_{g,t},l_{g,t},H_{t}\right)+\beta\mathbb{E}\left[V_{g,t+1}^{single}\left(\tilde{\mathbf{S}}_{g,t+1}\right)\right]+\tilde{u}\left(t\right),$$

where $\mathbf{a}_{g,t}^{div}=\left(c_{g,t},l_{g,t},A_{t}\right)$. Furthermore, even though we model (first time) divorce as an absorbing state, we do include an additional exogenously calibrated utility term, $\tilde{u}\left(t\right)$ to partly approximate remarriage possibilities. We thereby make the assumption that $\tilde{u}'(t)<0$, that is, this additional utility term is decreasing over time, mimicking the fact that remarriage probabilities are decreasing with age.

D.4 Retirement phase

We take a very simple approach for the retirement phase since this is not the focus of our paper. However, we need to include it for the empirical performance of our model. Without a retirement phase it would be very difficult to explain both labor supply profiles and homeownership rates, given that both are useful in the retirement phase as a source of savings to allow for sufficiciently high consumption after the labor active part of the lifecycle. We leave a better treatment of the retirement phase, and its impact on the value of marriage, as an important avenue for future research.

In retirement, individuals no longer work and no longer face earnings shocks. Instead of their labor income, they receive a pension, $y_{g}^{r},g\in\{M,F\}$. This is defined as a replacement rate, b^{r}, that is multiplied by the earnings in the last period in which the individual was working. Formally this means that preferences no longer include leisure for women. The state variables for $T_{r}\leq t\leq T_{d}$ are $\mathbf{S}_{t}=$

$(A_{t-1}, H_{t-1}, p_t, \theta_{M,t}, \theta_{F,t}, \tilde{\mu}_{M,t}, \tilde{\mu}_{F,t})$ and the choice variables are $\mathbf{a}_t = (c_{M,t}, c_{F,t}, A_t, H_t, D_t)$. The budget constraint then becomes

$$\begin{aligned} & A_t + x_t + \tilde{p}_t (H_t) + \Omega (H_t, H_{t-1}) \\ = & (1+R) A_{t-1} + y_M^r + y_F^r + y_t^{nl} + \tilde{p}_t (H_t) \mathbf{1} [H_{t-1} > 1], \end{aligned} \tag{15}$$

and the borrowing constraint

$$\begin{aligned} Debt_t &\leq \max\{Debt_{t-1}, \min\{LTV_t, LTI_t\}\} \\ LTV_t &= \lambda^H \tilde{p}_t (H_t), \\ LTI_t &= \lambda^Y (y_M^r + y_F^r). \end{aligned}$$

E Computational details for solving the model

Details on discretization. For housing prices, we use the Tauchen discretization method (Tauchen, 1986). We use 12 nodes for the house price grid. To approximate the wage and earnings processes, we need to take into account the age-dependent distribution generated through the presence of unit roots in the productivity shocks. To accommodate for this, we discretize the wage and earnings distributions at each age separately. We use 10 nodes for both female wages and male earnings. The grid for male earnings is also doubled to allow for the contingency of unemployment.

Model solution. We solve the model using backwards recursion. For each of the grid points, agents make decisions given that values for the next time period, $t+1$, are predetermined. We first solve the model for single individuals denoted as V_F^{single} and V_M^{single}. These values are then used to compute the value functions for married individuals. Solving the value function on each grid is slower than alternative methods (e.g., the endogenous grid point method), but the specific form of the borrowing constraint complicates its usage, given the potential for non-differentiabilities in the value functions. Given that the constraint set for (A_{t+1}, H_t) is not rectangular, and the constraint on H_t is endogenous on A_{t+1}, we adjust the state space following Bajari et al. (2013) such that debt levels satisfy the LTV-limit and the optimization runs over a rectangular space. Conditional on H_{t-1} and H_t, we then solve for consumption and female labor supply, subject to the debt-to-income limit and given these conditional choices, we then find the optimal values for H_t by selecting the level which yields the highest value. We use the *Nelder-Mead* algorithm to minimize the distance between the estimates of the auxiliary models on the observed and simulated data. We run 10,000 simulations of all the exogenous stochastic processes, incl. the joint earnings and wage process for individuals, as well as match qualities. To assure a global solution, we conducted multiple runs by varying the initial parameter values in our optimization procedure.

F Estimation details for wages and earnings

For the estimation of the variances for permanent productivity shocks we rely on the identification arguments in Meghir and Pistaferri (2004) and Blundell et al. (2008). In particular, let us define

$$\widetilde{\ln y}_{M,t} = \ln y_{M,t} - \alpha_M(t),$$

and

$$\widetilde{\ln w}_{F,t} = \ln w_{F,t} - \alpha_F(t).$$

This gives us the unexplained parts of earnings and wages as

$$\Delta\widetilde{\ln y}_{M,t} = \epsilon_{M,t},$$

and

$$\Delta\widetilde{\ln w}_{F,t} = \epsilon_{F,t}.$$

The variance of the innovations $\epsilon_{M,t}$ for men's earnings can then be identified using the single second moment

$$\sigma^2_{\epsilon_M} = Cov\left(\Delta\widetilde{\ln y}_{M,t}, \Delta\widetilde{\ln y}_{M,t-1} + \Delta\widetilde{\ln y}_{M,t} + \Delta\widetilde{\ln y}_{M,t+1}\right).$$

Similarly for the variance of innovations to women's wages:

$$\sigma^2_{\epsilon_F} = Cov\left(\Delta\widetilde{\ln w}_{F,t}, \Delta\widetilde{\ln w}_{F,t-1} + \Delta\widetilde{\ln w}_{F,t} + \Delta\widetilde{\ln w}_{F,t+1}\right).$$

These equations can be directly estimated from panel data on male earnings and female wages, after removing the age-profiles from the wage and earnings dynamics. We use a 2-step Heckman selection to estimate the wage process of women.

G Nonlabor income

We used linear projections on the state space to reduce the size of the state space. In particular, we ran regressions of (real) nonlabor income on age ($= t$) and either (real) wages $w_{F,t}$ (for women) or earnings (for men). We further controlled for educational attainment of respondents, as well as region and year FEs. From these we get values for α^{nl}.

H Further results from the estimated model

H.1 Matched moments

Table 16 shows that we fit the moments quite well, which implies that average outcomes in terms of the key household decisions such as (female) labor supply and housing demand are captured well by our structural model. We do note that we underestimate the crude divorce rates in our model through the latter period of the working phase, from 35 onwards.

Table 16: Model simulations and data.

Moments	*Data*	*Simulations*
White v. White: labor supply response		
△ Hours worked	-2.23	-1.81
Labor supply		
Hours worked married women 25-29	23	24
Hours worked married women 30-34	20	20
Hours worked married women 35-39	19	19
Hours worked married women 40-44	21	23
Hours worked married women 45-49	23	23
Hours worked single women 25-29	27	25
Hours worked single women 30-34	24	21
Hours worked single women 35-39	23	20
Hours worked single women 40-44	23	25
Hours worked single women 45-49	25	29
Employment rate married women 25-29	0.66	0.82
Employment rate married women 30-34	0.69	0.66
Employment rate married women 35-39	0.75	.0.67
Employment rate married women 40-44	0.80	0.80
Employment rate married women 45-49	0.83	0.85
Employment rate single women 25-29	0.75	0.86
Employment rate single women 30-34	0.73	0.72
Employment rate single women 35-39	0.75	0.69
Employment rate single women 40-44	0.80	0.85
Employment rate single women 45-49	0.84	0.98
Housing		
% of married individuals as homeowners by age 30	0.73	0.63
% of married individuals as homeowners by age 35	0.84	0.83
% of married individuals as homeowners by age 45	0.89	0.90
% of single men individuals as homeowners by age 30	0.67	0.55
% of single men individuals as homeowners by age 35	0.72	0.72
% of single men individuals as homeowners by age 45	0.72	0.9
% of single women individuals as homeowners by age 25	0.54	0.39
% of single women individuals as homeowners by age 35	0.53	0.59
% of single women individuals as homeowners by age 45	0.65	0.78
Median housing value married couples 30 - 49	4.62	4.85
Median housing value single men 25 - 49	3.79	3.19
Median housing value single women 25 - 49	3.96	3.20
White v. White: crude divorce rates.		
Divorce rate 25-49	0.012	0.012
Crude divorce rates		
Divorce rate 25-29	0.029	0.0205
Divorce rate 30-34	0.027	0.0307
Divorce rate 35-39	0.024	0.0106
Divorce rate 40-44	0.020	0.0055
Divorce rate 45-49	0.0155	0.002

I Details on computation of welfare costs

We use a consumption-equivalent variation measure to calculate welfare costs of a change in the LTI-limit (see, e.g., Fehr and Kindermann, 2018). To do so we first fix a set of simulated paths for match quality, male earnings, female wages, and house price shocks. Then, for a given household we can compute the resulting (simulated) paths for the outcome variables $c_{M_t}, c_{F,t}, x_t, H_t, l_{F,t}$. With these we can compute, for each individual within each household their expected (lifetime) utility, as follows:

$$U_g = \sum_{t=1}^{T_d} \beta^{t-1} u_g \left(c_{g,t}, l_{g,t}, H_t, \theta_{g,t} \right).$$

Similarly, the expected lifetime utility can also be computed in the counterfactual case. That is, for the new house prices simulate new consumption, leisure and housing demand paths $\tilde{c}_{M,t}, \tilde{c}_{F,t}, \tilde{x}_t, \tilde{H}_t, \tilde{l}_{F,t}$ (for the same set of simulated paths of match quality, male earnings and female wages). This gives us the expected (lifetime) utility in the counterfactual situation:

$$\tilde{U}_g = \sum_{t=1}^{T_d} \beta^{t-1} u_g \left(\tilde{c}_{g,t}, \tilde{l}_{g,t}, \tilde{H}_t, \theta_{g,t} \right).$$

To evaluate the impact on individual welfare, we then find a value z such that:

$$\sum_{t=1}^{T_d} \beta^{t-1} u_g \left((1+z) c_{g,t}, l_{g,t}, H_t, \theta_{g,t} \right) = \tilde{U}_g.$$

We also calculate the Hicksian equivalent variation (HEV) which can be calculated in a very similar way, in particular we then find a value z^{HEV} that solves the following:

$$\sum_{t=1}^{T_d} \beta^{t-1} u_g \left(\left(1+z^{HEV}\right) c_{g,t}, \left(1+z^{HEV}\right) l_{g,t}, H_t, \theta_{g,t} \right) = \tilde{U}_g.$$

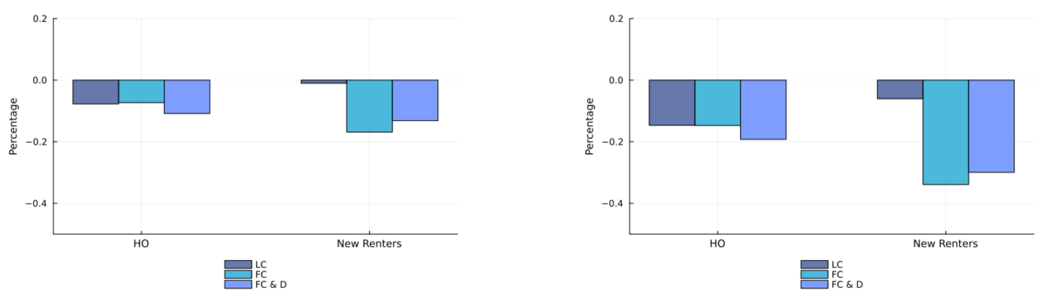
That is, the HEV calculates the proportional change in both private consumption and leisure so as to make the individual indifferent again between the baseline and counterfactual economy.

J Further results welfare cost LTI tightening

We calculated the welfare costs in two alternative ways. First, in contrast to the main paper where we checked the consumption equivalent change to make an individual indifferent between the baseline economy and the counterfactual scenario, we also looked at just the consumption and leisure equivalent change. This has the added benefit of being more standard. Second, we also calculated the consumption equivalent change as a welfare measure where we further keep women's labor supply fixed at their choices in the baseline scenario.

The results are presented in Figure 8. The results are qualitatively quite similar to those in the main paper. In particular, married women seem to be much better insured to the LTI tightening when changes in intra-household bargaining are allowed. This suggests that the welfare cost when ignoring this channel might be overestimated. Quantitatively, there is a slight increase in the welfare cost incurred by women compared to those presented in the main paper, which is the result of not being able to receive enough compensation through the labor supply channel.

Figure 8: *Welfare costs of LTI-tightening for married women.*



(a) *Hicksian equivalent variation* (b) *Fixed labor supply*

Notes: own calculations based on the same set of 10,000 simulations.