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HAS US HOUSEHOLD DELEVERAGING ENDED? A MODEL-BASED ESTIMATE OF EQUILIBRIUM DEBT

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In 2014 all ECB publications feature a motif taken from the €20 banknote.

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Abstract

The balance sheet adjustment in the household sector was a prominent feature of the Great Recession that is widely believed to have held back the cyclical recovery of the US economy. A key question for the US outlook is therefore whether household deleveraging has ended or whether further adjustment is needed. The novelty of this paper is to estimate a time-varying equilibrium household debt-to-income ratio determined by economic fundamentals to examine this question. The paper uses state-level data for household debt from the FRBNY Consumer Credit Panel over the period 1999Q1 to 2012Q4 and employs the Pooled Mean Group (PMG) estimator developed by Pesaran et al. (1999), adjusted for cross-section dependence. The results support the view that, despite significant progress in household balance sheet repair, household deleveraging still had some way to go as of 2012Q4, as the actual debt-to-income-ratio continued to exceed its estimated equilibrium. The baseline conclusions are rather robust to a set of alternative specifications. Going forward, our model suggests that part of this debt gap could, however, be closed by improving economic conditions rather than only by further declines in actual debt. Nevertheless, the normalisation of the monetary policy stance may imply challenges for the deleveraging process by making a given level of household debt less affordable and therefore less sustainable.

Keywords: Household deleveraging, Equilibrium debt, Pooled Mean Group estimator, Heterogeneous dynamic panels.
JEL Classification: C13, C23, C52, D14, H31

Non-technical summary

Household deleveraging in the US has acted as significant headwind to consumption and activity in recent years, holding back the recovery. Despite the substantial balance sheet adjustment that has resulted from both the paying-down of debt and defaults, a key question with regard to the US outlook is whether deleveraging has ended or whether further adjustment is needed. The challenge in answering this question comes from the fact that not only there is no obvious benchmark to which the debt ratio should converge, but also because the level of sustainable debt is likely to be time-varying and evolve in line with economic fundamentals. In this paper, we aim at providing time-varying, model-based estimates of the household equilibrium debt-to-income ratio in order to track progress in household deleveraging needs. Our approach is based on a panel error correction framework using US state-level data over the period 1999Q1-2012Q4. The main results are as follows:

1. Since around 2002-03, the US household debt-to-income ratio began to deviate increasingly from its estimated equilibrium level. The difference between the actual and the equilibrium ratio – the so-called debt gap – increased substantially between that time and the end of 2008, due to: (i) the actual debt-to-income ratio rising at a faster pace than what was suggested by equilibrium debt, and (ii) since mid-2007, the equilibrium ratio starting to decline on account of deteriorating fundamentals, such as lower house prices, higher uncertainty, more pessimistic income expectations and reduced collateral availability;

2. The debt gap has begun to shrink since 2009, due to deleveraging undertaken by households and, more recently, also due to rising equilibrium debt on account of improved fundamentals and supportive low interest rates;

3. The actual debt-to-income ratio continued to exceed its estimated equilibrium by around 6 percentage points as of 2012Q4, implying that although significant progress has been made (closing the debt gap by around 80% relative to its peak), household balance sheet repair still had some way to go;

4. The number of states in need of adjustment has been shrinking: virtually all US states had a debt gap above zero when the national debt gap was at its peak (2008Q4), but as of 2012Q4 the debt adjustment seemed to have been completed in one-third of the states. States with pronounced boom-bust cycles in their housing markets (Arizona, California, Florida and Nevada) have managed to achieve substantial balance sheet repair, but further adjustment appeared to lie ahead for California and, to a much lesser extent, Nevada;

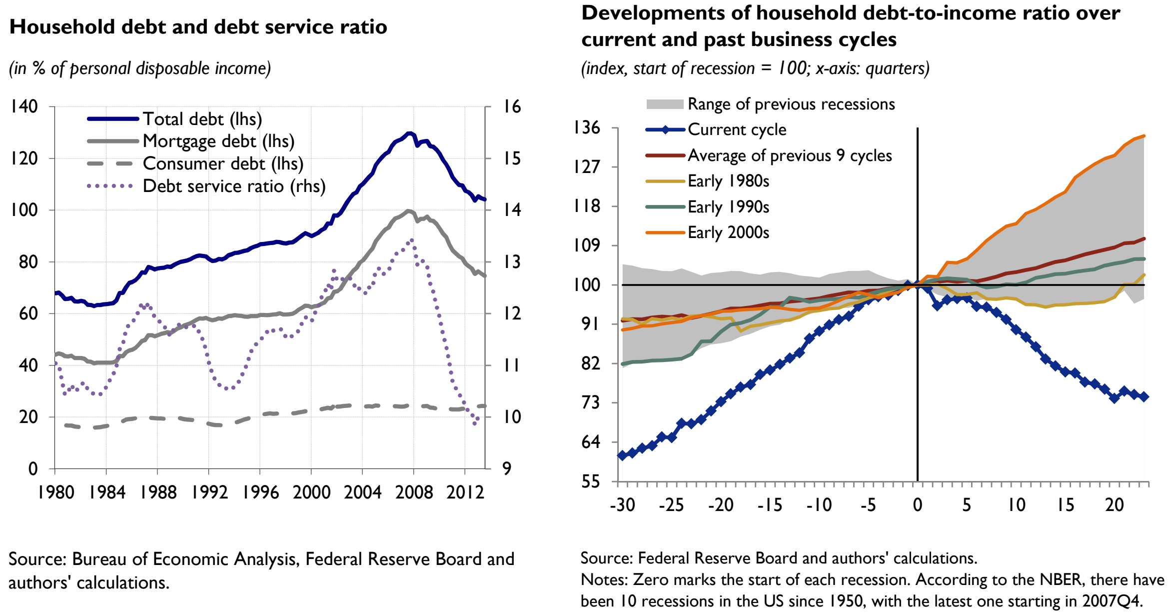
5. Looking ahead – should economic conditions continue to improve as expected – around half of the remaining debt gap could be closed by end-2015 via a rise in the sustainable debt-to-income ratio, leaving the remaining half to be corrected via further deleveraging carried out by households;

6. On the other hand, potentially higher mortgage interest rates, reflecting the normalisation of monetary policy and exit from non-standard measures, will imply larger adjustment needs in the future by pushing down, *ceteris paribus*, the sustainable debt-to-income ratio.

1 Introduction

The balance sheet adjustment in the household sector has been a prominent feature of the most recent US recession and subsequent recovery. The beginning of the economic downturn in late-2007 broadly coincided with the start of a sustained reduction in household liabilities relative to income – or household deleveraging – which contrasted with the strong build-up of debt before the crisis. From a peak of around 130% in 2007Q4, the household debt-to-income ratio fell by more than 25 percentage points to around 104% in 2013Q3, led by sustained declines in mortgage debt (left-hand panel of Figure 1).[1] Forecasters and economists broadly agree that household deleveraging has acted as an important drag on the recovery.[2] However, a key question for the US outlook – how much further deleveraging is needed – remains unanswered, as there is no obvious benchmark to which the debt ratio should converge. History appears to offer little guidance as regards the adjustment needs in the current cycle, as the debt level at the start of the recent recession was unprecedented and recent swings in the household debt-to-income ratio are unusual by the standards of previous recessions (right-hand panel of Figure 1).

Figure 1: Household debt, service ratio and household debt-to-income over the business cycles



Historically, the ratio increased modestly prior to recessions (on average by 8 percentage points in the 30 quarters preceding a recession). This compares with a much sharper rise of 39 percentage points in the run-up to the most recent recession. Moreover, when a recession ends, households typically start to build up debt again, reflecting an easing in credit standards and an increase in credit availability, combined with rising confidence and an upward shift in future income expectations, which support credit demand. For example, Schmitt (2000) finds that rising consumer indebtedness is a normal occurrence in an economic expansion. The current

[1] A significant proportion of the reduction in debt has resulted from defaults by households, with estimates varying from around 40% (Brown et al. 2010) to 70% (McKinsey Global Institute 2012).

[2] See for example Mian and Sufi (2010) and Mian et al. (2013) on the link between household leverage and activity, and Eggertsson and Krugman (2012), and Koo (2008, 2013) on the notion of "deleveraging crisis" and "balance sheet recessions".

cycle, however, has not shared this feature, with the debt-to-income ratio continuing to decline even in the fourth year of the economic recovery. The economic fundamentals have been less supportive of a renewed upturn in the credit cycle as credit standards have remained relatively tight, income expectations and the labour market have improved only gradually, and house prices have recovered only slightly from the sharpest correction since the Great Depression.

This paper proposes a novel approach to examine the question of how far indebtedness stands from its sustainable level at any point in time by estimating a time-varying equilibrium household debt-to-income ratio determined by economic fundamentals. This approach allows us to assess whether household indebtedness moved beyond what was suggested by its fundamentals during the recent credit boom, as well as to track progress in household deleveraging in the current phase of balance sheet adjustment. The paper uses aggregated state-level data from the Federal Reserve Bank of New York's (FRBNY) Consumer Credit Panel, a nationally representative sample drawn from anonymised Equifax credit data, over the period 1999Q1 to 2012Q4 and employs the Pooled Mean Group (PMG) estimator developed by Pesaran et al. (1999), adjusted for cross-section dependence. It fills a gap in the literature, as existing studies that evaluate the sustainability of household debt commonly use as benchmarks either pre-crisis trends or historical episodes in other countries that experienced strong debt cycles (see (McKinsey Global Institute 2012). The factors mentioned above, however, would indicate that the equilibrium debt-to-income ratio is likely to be time-varying and to evolve with shifts in expectations, the demographic structure and the availability and cost of credit. This suggests that extrapolating historical (linear) trends to estimate equilibrium debt-to-income ratios may be inappropriate for assessing the extent to which the household debt burden is excessive at any point in time.

Overall, while debt sustainability for the public sector has been researched extensively (see for example Wyplosz 2007), the empirical literature on sustainable household debt ratios is rather scarce. A number of empirical studies, such as Fernandez-Corugedo and Muellbauer (2006), Dynan and Kohn (2007), and Aron and Muellbauer (2013) study the fundamental determinants of household debt without, however, providing quantitative conclusions regarding its sustainability. More recently, a study by Cuerpo et al. (2013) proposes two approaches to estimate sustainable debt against which to assess the actual indebtedness in the non-financial private sector in EU Member States. The first one is based on the "threshold approach", a static measure computed as the upper quartile of the distribution of debt-to-GDP ratio. The second one, the "stationarity approach", is a time-varying measure that relies on the notion of stationarity of household debt in terms of notional leverage, i.e. when household debt evolves in line with total deflated assets. However, this modelling approach ignores the possibility that households' sustainable debt ratio could also depend on factors other than assets, such as income expectations, uncertainty, the cost and access to credit.

Regarding the theoretical literature, the contribution by Barnes and Young (2003) is closely related to our work. The authors employ a calibrated partial equilibrium overlapping generations model for US household debt-to-income, which can be used to assess the causes of the change in aggregate debt based on a number of fundamental drivers (including real interest rates, income growth expectations and demographic changes). The difference between the behaviour implied by their model and actual developments may also be indicative of an unsustainable disequilibrium.

Our approach also builds on a number of empirical studies that have employed the PMG estimator for estimating equilibria and the resulting gaps for other economic variables. For example, Kiss et al. (2006) employ a PMG approach to assess whether strong credit growth to the private sector in a number of Central and Eastern European countries reflects equilibrium convergence or excessive credit. More recently, Poghosyan (2012) studies the determinants of sovereign bond yields in advanced economies with a PMG approach. Along the same lines, Csonto and Ivaschenko (2013) study the determinants of sovereign bond spreads for 18 emerging market economies, analysing, in particular, whether bond prices in these countries were in line with fundamentals at any given point in time.

The contribution by Holly et al. (2010) is also important for our study regarding the econometric framework. They use the PMG and the common correlated effects mean group methodology to estimate a panel error correction model of house prices for US states. The authors find that real house prices have been rising broadly in line with fundamentals up to the end of their sample period (2003), although there have been a number of outlier states (including California, New York and Massachusetts).

Our results show that the debt-to-income ratio in the US household sector was broadly in line with what was suggested by equilibrium debt up to around 2002-03. Later on, a positive debt gap started to emerge as actual debt rose at a faster pace than equilibrium debt. Since mid-2007, the widening in the debt gap was reinforced by a decline in equilibrium debt, reflecting deteriorating fundamentals, such as lower house prices, higher uncertainty, more pessimistic income expectations and reduced collateral availability. These factors were partially offset by lower mortgage rates. At the time when the national debt gap was at its peak (2008Q4), all the US states had a debt gap above zero, although with different deleveraging needs. The deleveraging process, initiated in 2009, allowed the debt gap to be closed by roughly 80% from its peak until the end of 2012. Our baseline estimates indicate a remaining gap of around 6 percentage points by the end of 2012, with households in one-third of the states no longer facing deleveraging needs. We also find that the conclusions reached are rather robust to alternative specifications.

The remainder of the paper is organised as follows. In the next section, we introduce the econometric framework used to estimate the equilibrium debt-to-income ratio, while in Section 3, we report the estimated coefficients for our baseline model and present the resulting equilibrium debt ratios and the implied debt gaps at the aggregate level. In Section 4, we examine the robustness of our results. In turn, in Section 5 we focus on the heterogeneity of household deleveraging at the state level. Section 6 concludes.

2 Modelling the household debt-to-income ratio in the long-run

2.1 Econometric framework

In order to assess whether household indebtedness has moved beyond what is justified by its fundamentals during the recent credit boom and to track progress in household deleveraging in the current phase of balance-sheet adjustment, we propose an empirical approach to estimate a time-varying equilibrium debt-to-income ratio. We model the ratio of US household debt to income in a panel error correction framework for the 50 US states (plus the District of Columbia) over the period 1999Q1 to 2012Q4, using aggregated, state-level summary data from the FRBNY Consumer Credit Panel, a nationally representative sample drawn from anonymised Equifax credit data.[3] We estimate our model with the PMG estimator developed by Pesaran et al. (1999), which assumes identical long-run coefficients across states, but allows for a differentiated response to short-term factors depending on state-specific characteristics. More specifically, we take as a staring point the following equation (see Appendix A for the derivation):

$$\Delta d_{it} = \mu_i + \phi_i(d_{i,t-1} - \theta_1' X_{it}) + \delta_{i1}^{*\prime} \Delta X_{it} + u_{it} \tag{1}$$

$$\text{where} \quad X_{it} = \begin{bmatrix} hp_{it} \\ hown_{it} \\ i_{it} \\ ur_{it} \\ dem_{it} \\ ltv_{it} \\ frcl_{it} \end{bmatrix}, \ \theta_1 = \begin{bmatrix} \theta_{11} \\ . \\ . \\ . \\ . \\ . \\ \theta_{17} \end{bmatrix} \ \text{and} \ \delta_{i1}^* = \begin{bmatrix} \delta_{i11}^* \\ . \\ . \\ . \\ . \\ . \\ \delta_{i17}^* \end{bmatrix}$$

where d is the household debt-to-income ratio derived from aggregated, state-level summary data, X is a set of explanatory variables (which will be introduced in the next section), θ and δ^* are respectively the long-run and the short-run coefficients, and ϕ is the speed of adjustment with an expected negative sign. Over the long term a stable relationship between the debt-to-income ratio and the explanatory variables would be expected, while allowing for temporary deviations from this relationship. The gap between the actual debt-to-income ratio and its estimated equilibrium determined by the long-run relationship would be interpreted as deviations from "sustainable" levels.

The standard PMG estimator implicitly assumes that the errors u_{it} of Equation (1) are independently distributed across states. However, if this is not the case, the presence of cross-section dependence implies that unobserved factors in the error term could be correlated with the explanatory variables (Pesaran 2006). Neglecting this econometric issue could lead to biased estimates and to spurious inference. Testing for the presence of cross-section dependence using a test by Pesaran (2004) suggests that the assumption of independence might not be valid, which implies the presence of unobserved common factors in the error term (possibly related to the housing bubble or the financial crisis). As a result, we employ a modified PMG estimator that has been developed to allow for correlation across panel members due to unobserved common time-specific effects: the Common Correlated Effects Pooled Mean Group (CCEPMG) estimator – see Pesaran (2006), Binder and Offermanns (2007), Holly et al. (2010), and Chudik and

[3]Household debt as reported in the Consumer Credit Panel (CCP) draws on detailed anonymous Equifax credit-report data and as such differs somewhat from the Board of Governors' Flow of Funds (FoF) Accounts measure (shown in Figure 1). The FRBNY's CCP measure of household debt to personal disposable income is on average 15 percentage points lower than the Flow of Funds measure (1999-2012), but shares broadly the same dynamics over time (correlation between the two series: 98.2%) and the major components – home mortgage debt and consumer credit debt – are directly comparable. In our empirical analysis, we scale the FRBNY's debt measure to personal income.

Pesaran (2013).[4] To allow for cross-section dependence, the error term in the case of CCEPMG is specified as:

$$u_{it} = \lambda_i^{'} f_t + \varepsilon_{it} \tag{2}$$

where an unspecified number of unobserved common factors f_t with idiosyncratic factor loadings λ_i capture time-variant heterogeneity and cross-section dependence, while ε_{it} are now idiosyncratic errors independently distributed across i and t. The CCEPMG estimator – our preferred specification – is constructed by augmenting the standard PMG with cross-section averages of the variables as additional regressors, in order to account for the unobserved factors (see Appendix A for more details):

$$\Delta d_{it} = \mu_i + \phi_i(d_{i,t-1} - \theta_1' X_{it}) + \delta_{i1}^{*\prime} \Delta X_{it} + \alpha_i \overline{d}_t + \beta_i \overline{X}_t + \gamma_i \Delta \overline{d}_t + \eta_i \Delta \overline{X}_t + \varepsilon_{it} \tag{3}$$

where $\overline{d}_t$ and $\overline{X}_t$ are averages of the dependent variable and the regressors across states, computed at every time period t. Equation (3) is estimated with one lag (in levels) for all the variables, as suggested by the Bayesian Information Criterion (see Table C.1 in Appendix C).

2.2 Data

We employ the household debt-to-income ratio as dependent variable, rather than the debt-to-asset ratio for the following reasons. First and foremost, while both measures are plausible and emphasise different aspects of indebtedness, data availability prevents us from using financial assets or the debt-to-asset ratio: while income is available at the state level, financial assets (and more generally household assets) are not. Secondly, using the debt-to-assets ratio might be problematic during asset price booms and busts. In fact, the denominator (assets) rises during a boom, which might mask the "true" build-up/misalignment of debt until the asset price bubble bursts. This was evident in both the dotcom bubble in the late 90s and the housing bubble in the mid-2000s. Thirdly, our focus is on identifying potentially unsustainable developments in household debt, which is better captured by the debt-to-income ratio as it emphasises debt servicing capacity. Moreover, financial assets are highly concentrated within the upper percentiles of the income distribution, more so than debt. Hence, there are issues related to distributional aspects when using aggregate (not micro) data, since those households who hold most of the financial assets do not necessarily coincide with those who hold the bulk of debt.[5] Finally, the debt-to-income ratio is often the choice in the related economic literature, such as in Dynan and Kohn (2007), Mian and Sufi (2010), Aron and Muellbauer (2013), McKinsey Global Institute (2012), Mian et al. (2013), and several other studies by regional Federal Reserve Banks in the US.

Our choice of explanatory variables for determining the ratio of household debt-to-income follows broadly previous studies by Barnes and Young (2003), Fernandez-Corugedo and Muellbauer

[4] An alternative way of dealing with cross-section dependence is to use the Augmented Mean Group (AMG) estimator introduced in Eberhardt and Teal (2010), which we also report in Section 4.4.

[5] In the discussion in Dynan (2012), C. Carroll notes that the debt-to-income ratio seems preferable to the debt-to-asset ratio because a large fraction of households have few assets aside from their home.

(2006), and Dynan and Kohn (2007). In particular, we consider the following variables at the state level (in parenthesis the expected sign of the long-run effect is shown) – see Appendix B for the sources and descriptive statistics:

- **Logarithm of the FHFA house price index-to-income ratio (*hp*)**: increases in house prices will lead to higher borrowing via raising desired consumption through the traditional wealth effect. Higher house prices also increase household debt levels via collateral effects. (+)

- **Homeownership rate (*hown*)**: we use the homeownership rate to proxy for collateral effects beyond the effects of the pricing of collateral, which may be driven by unsustainable trends. An increase in the homeownership rate should be associated with a higher debt-to-income ratio, as the higher stock of housing assets held by households would reduce borrowing constraints and raise available collateral for borrowing.[6] (+)

- **Nominal interest rate on conventional mortgages (*i*)**: higher real interest rates reduce the affordability of household debt and make saving more attractive. Moreover, for a given real rate, a higher nominal interest rate is likely to increase the necessary nominal interest payments relative to income on a prospective loan, making it more likely to fall above some upper bound imposed by the lender. (-)

- **Unemployment rate (*ur*)**: as suggested by Fernandez-Corugedo and Muellbauer (2006), we use the unemployment rate to proxy for both income expectations and uncertainty. Higher future income expectations should lead to increased current consumption due to consumption smoothing, and thus to higher debt. Moreover, lower uncertainty would imply less need for precautionary reserves, which tends to increase borrowing. (-)

- **35-54 age group (*dem*)**: the demographic structure of the population also plays a role. According to the life-cycle model, households incur debt when they are young, save when they are older and dis-save towards the end of their life. As shown by Emmons and Noeth (2013), household mortgage debt tends to be highest for the late 30s to early 40s age group. We use the share of the population in the 35-54 age group, in line with the findings by Barnes and Young (2003) that this age group accounted for the largest increase in US household debt from 1989 to 1998. (+)

- **Loan-to-value ratio for first-time borrowers (*ltv*)**: financial innovation can lead to greater risk-sharing and increased credit supply, leading to more relaxed credit standards and improved access to credit for households. Theoretical considerations suggest that lower down payment requirements to undertake a mortgage would lead to higher aggregate debt by relaxing borrowing constraints. We use the loan-to-value ratio (available only at the US national level) to proxy for financial innovation and credit availability. (+)

- **Foreclosure rate (*frcl*)**: An increase in default rates and foreclosures may reduce the stock of outstanding debt by increasing the amount of non-performing loans being written off by lenders (see Brown et al. 2010). Unlike the other variables, we allow the foreclosure

[6] Borrowing constraints of this type, where households can borrow up to a fraction of available collateral – usually proxied by real estate holdings – is fairly standard in the macroeconomic literature. See for example Kiyotaki and Moore (1997) and Iacoviello (2004).

rate to affect debt only in the short run (i.e. θ_{17}=0 but $\delta^*_{17} \neq$0) as we do not find compelling reasons for it to have an impact on long-run equilibrium debt. (-)

2.3 Panel unit root and cointegration tests

Before turning to the estimation, we conduct a number of preliminary tests on the data. First, we test whether the variables are stationary by running a set of (panel) unit root tests (Tables C.2 and C.3 in Appendix C). The results broadly support the unit root hypothesis for the variables over our sample period. Second, having established that the variables are non-stationary, we test whether there is a stable long-run cointegrating relationship between the variables that would validate our approach of modelling the relationship within a cointegrating framework. We use three methods to check for cointegration.

The first one is a panel cointegration test by Gengenbach et al. (2009) between the debt-to-income ratio and the explanatory variables available at the state level.[7] The advantage of this test is that it accounts for cross-section dependence, which is suitable in our case. The results, shown in Table C.4 in Appendix C, suggest the presence of cointegration among our variables. In the second method, we interpret the negative sign and significance of the error-correction parameter ϕ as an indicator of cointegration between the variables, as often done in the literature (see Egert et al. 2004, Kiss et al. 2006, Durdu et al. 2013, Poghosyan 2012). Indeed, this turns out to be the case as we will show in Section 3.1. Finally, we conduct panel unit root tests on the state-level debt gaps, i.e. the difference between actual debt-to-income ratio and the estimated long-run equilibrium level, which is equivalent to a unit root test on the equilibrium residuals. Again, as we will show later in Section 3.1, the results suggest that the residuals are stationary, thus supporting the validity of our cointegration approach.

3 Estimation results

3.1 Baseline specification

Table 1 reports the estimation results of our preferred CCEPMG specification. Focusing on the long-term coefficients, all the variables are statistically significant and with the expected sign. House prices are statistically significant and with the expected positive sign throughout all specifications, suggesting that a 1% increase in the house price-to-income ratio would lead to an increase in the household debt-to-income ratio by a range between 0.2 and 1.0 percentage points. The homeownership rate is also positive and statistically significant in the last two columns, supporting the view that states with higher homeownership rates have in general higher debt-to-income ratios, as a larger stock of housing assets serves as collateral for borrowing. In turn, our estimations suggest that higher interest rates create a disincentive for households to take on more debt, in line with economic theory; a one-percentage point increase in nominal interest rates would lead to a decline in the household debt ratio of about 0.6 to 5.1 percentage points. Higher uncertainty and lower future income expectations, as proxied by the unemployment rate,

[7] We implement the Gengenbach et al. (2009) cointegration test based on a modified Stata code available from M. Eberhardt's website at: https://sites.google.com/site/medevecon

exert downward pressure on the debt ratio, as expected. The positive sign of the coefficient on the demographics variable is in line with the findings in the literature suggesting that a larger proportion of the early 30s to late 40s age group in the population is associated with higher household debt. Finally, the loan-to-value ratio, a proxy for credit supply, is positive and statistically significant, confirming the theoretical prediction that higher credit availability allows households to take on more debt.

Table 1: Estimation results based on the CCEPMG specification

	(1)	(2)	(3)	(4)	(5)	(6)
Long-run coefficients						
House prices	104.245***	100.144***	58.138***	54.729***	24.051***	24.320***
	(9.496)	(10.130)	(6.737)	(5.107)	(5.031)	(4.818)
Homeownership rate		-0.041	0.097	0.112	0.325***	0.244***
		(0.362)	(0.202)	(0.141)	(0.081)	(0.079)
Interest rates			-3.784***	-5.110***	-0.552*	-1.727***
			(0.326)	(0.373)	(0.311)	(0.370)
Unemployment rate				-2.427***	-1.270***	-1.277***
				(0.545)	(0.346)	(0.349)
35-54 age group					3.534***	3.386***
					(1.130)	(1.043)
Loan-to-value ratio						0.526***
						(0.087)
Speed of Adjustment	-0.090***	-0.088***	-0.134***	-0.198***	-0.385***	-0.378***
	(0.010)	(0.010)	(0.013)	(0.020)	(0.022)	(0.023)
Short-run coefficients						
Δ House prices	32.444***	32.401***	33.525***	28.419***	28.177***	28.495***
	(4.691)	(4.760)	(4.788)	(5.180)	(5.251)	(5.132)
Constant	-10.057***	-9.807***	-9.814***	-8.060***	35.044***	4.09
	(1.304)	(1.317)	(1.403)	(2.522)	(8.143)	(7.686)
Half-life	7.3	7.5	4.8	3.1	1.4	1.5
Observations	2,805	2,805	2,805	2,805	2,805	2,805
Hausman test (p-value)	0.915	0.899	0.729	0.002	0.537	0.513

Notes: CCEPMG estimates corresponding to the common correlated effects specification of the PMG. The dependent variable is total household debt-to-income ratio. We include the following cross-section averages: house price-to-income ratio, unemployment rate, and 35-54 age group (see Appendix A). The lag structure (1 lag) was selected using the Schwartz Bayesian criterion. Foreclosures can only influence household debt in the short run. Due to space limitation, most of the short-run coefficients are omitted, with the exception of house prices. Standard errors are shown in parentheses. Asterisks, *, **, ***, denote, respectively, statistical significance at the 10, 5 and 1% levels. The half-life estimates indicate the number of quarters it takes to halve the gap between actual and equilibrium debt-to-income ratio. The Hausman test reports p-value under the null hypothesis that the CCEPMG estimator is both efficient and consistent, i.e. that the long-run homogeneity restriction is valid.

The speed of adjustment, or the error-correction coefficient, is negative and statistically significant throughout all specifications, supporting the cointegration hypothesis between the household debt-to-income ratio and the set of determinants included in the model. In particular, column (6) suggests that the average speed of adjustment of 0.378 means that the gap between actual and equilibrium debt would be closed relatively fast, implying a half-life of 1.5 quarters.[8] Furthermore, we apply panel unit root tests on the state-level equilibrium residuals (the so-

[8] One can expect half of the gap to be closed in line with the half-life estimate only in the case of a single shock. In the presence of multiple shocks over time, actual debt might deviate from its estimated equilibrium for longer than suggested by the half-life.

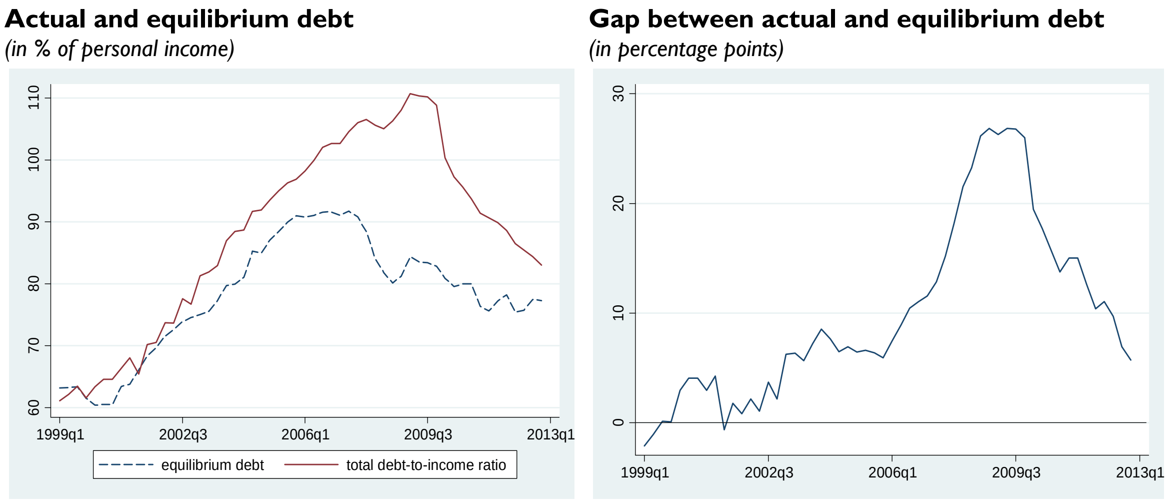
called debt gaps). As shown in Table C.5 in Appendix C, the results suggest that the residuals are stationary, supporting the validity of our cointegration approach. In terms of the model specification, the Hausman test does not reject the homogeneity restriction of the CCEPMG estimator, which assumes that the long-run coefficients are the same across the US states.[9]

3.2 Equilibrium debt and debt gaps over 1999-2012

To estimate the equilibrium debt-to-income ratio (and the implied debt gap) from our preferred CCEPMG model at the national level, we take the long-run coefficients, assumed to be homogeneous across states, and the long-run aggregate constant term (see Equation 8 in Appendix A on the aggregation method for equilibrium debt). To compute the aggregate long-term constant we have taken into account the fact that the short-term constants and the coefficients on the lagged dependent variables are not necessarily independently distributed across panel members, as shown in Equation 9 in Appendix A.

Our estimates suggest that until around 2006, equilibrium debt was increasing mainly on account of the observed strong rise in house prices, supportive interest rates and financial liberalisation, implying higher credit availability. Looking at the misalignment between actual and equilibrium debt, Figure 2 shows that the actual debt-to-income ratio was broadly in line with fundamentals up until the end of 2002, with the latter reflecting particularly the observed strong rise in house prices and supportive interest rates.

Figure 2: Actual and equilibrium debt-to-income ratio and implied gap



Source: FRBNY/Equifax Consumer Credit Panel and authors' calculations.
Notes: Last observation refers to 2012Q4.

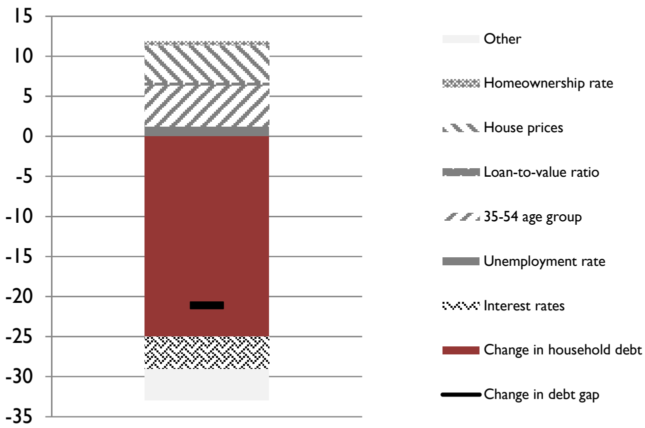
After 2002, however, the debt gap started to open up at the national level on account of a faster

[9]One of the advantages of using state-level data relates to the efficiency gain from pooling the panel data when employing the CCEPMG, thus making use of the richness contained in the cross-section dimension, while at the same time capturing the prevailing heterogeneity across US states. This efficiency gain becomes apparent when estimating the same error-correction specification on US aggregated data – practically meaning that we are treating the US as a single state – where none of the variables turns out to be statistically significant. Moreover, in aggregate time series analysis it is more difficult to identify the effects on the ratio of debt to income of particular variables which move slowly over time, such as demographic variables. These effects could thus be better captured through the cross-section dimension of the panel data.

increase in actual debt relative to what was suggested by equilibrium debt. Since 2007, the increase in the debt gap was reinforced by a decline in the equilibrium debt-to-income ratio, reflecting deteriorating fundamentals, such as lower house prices, unfavourable demographics, as well as higher uncertainty, more pessimistic income expectations and reduced collateral availability (Figure C.1 in Appendix C). The deterioration in the fundamentals was partially offset via declines in the mortgage rate which prevented an even sharper fall in equilibrium debt.

The debt gap reached its peak in late 2008, which broadly coincided with the peak in the actual debt-to-income ratio.[10] Thereafter, the gap began to shrink due to stronger deleveraging undertaken by households that outweighed the decline in the equilibrium debt ratio. More recently, in the course of 2012 the gap has shrunk not only due to the on-going household deleveraging, but also reflecting a gradual stabilisation and rise in the equilibrium debt-to-income ratio, on account of improved fundamentals and supportive low interest rates. In particular, lower mortgage rates have supported progress towards closing the debt gap since 2009, by increasing the level of the equilibrium debt-to-income ratio by 4 percentage points (Figure 3). This means that part of the relief from reduced balance sheet distress in the household sector is, to some extent, policy-induced, driven by the exceptional monetary accommodation via near-zero interest rates and purchases of mortgage securities under Quantitative Easing (QE).

Figure 3: Contributions to changes in the debt gap since its peak in 2008Q4 until 2012Q4
(in percentage points)



Source: Authors' calculations.
Notes: The components in black-and-white texture drive the changes in equilibrium debt, which is one determinant of the debt gap (the other being changes in actual household debt, shown in red).

Focusing on the end-point, the actual debt-to-income ratio at the national level in 2012Q4 exceeded its estimated equilibrium value, suggesting that household deleveraging still had some way to go before closing the gap. The estimated debt gap stood at almost 6 percentage points as of 2012Q4, which compares with a gap at the peak of around 27 percentage points in 2008Q4, meaning that almost 80% of the deleveraging process had been already completed.

As a robustness check, we compute the equilibrium debt-to-income ratio for the US aggregate from the aggregation of state-level equilibrium estimates, derived by either taking variable or fixed weights (the 1999-2012 average) for income. This is an alternative method to the one used throughout the paper, i.e. by applying the long-run coefficients and the long-run aggre-

[10]The peak in the debt-to-income ratio as measured by the FRBNY occurred in 2009Q1.

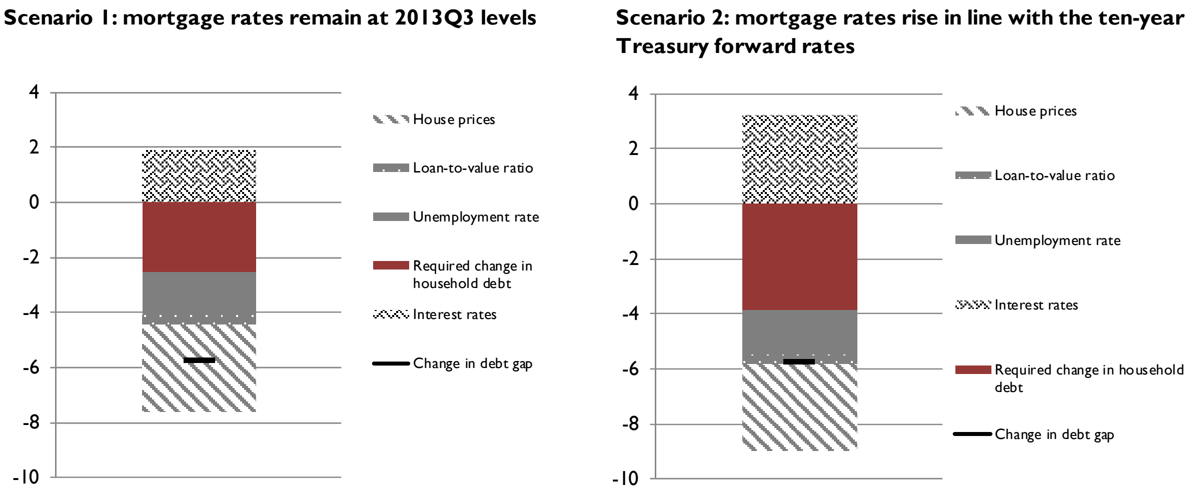
gate constant term directly to the available national data. Overall, our baseline specification, the CCEPMG, remains robust to this alternative aggregation method. Specifically, using this bottom-up approach, equilibrium debt would stand slightly above the baseline specification in the period prior to the crisis (implying a somewhat smaller gap), but afterwards it would overlap with the CCEPMG (see Figure C.3 in Appendix C for the version with variable weights).

3.3 Scenarios for closing the debt gap in 2013-15

After computing equilibrium debt at the national level for the 1999-2012 period, in this section we build some illustrative scenarios for how the remaining debt gap, as of 2012Q4, could be closed over 2013-15.

The time-varying nature of equilibrium debt implies that the gap could be closed by a combination of a further decline in the actual debt ratio and a continued rise in the equilibrium debt ratio. An improvement in economic conditions, such as a decline in unemployment or rising house prices, would cause the equilibrium debt ratio to rise, implying smaller adjustment needs. For instance, assuming that mortgage rates remain at 2013Q3 levels, the unemployment rate declines from 7.8% in 2012Q4 to 6.2% in 2015Q4 and the house price-to-income ratio rises gradually by 8.2 percentage points between 2012Q4 and 2015Q4 (in line with Macroeconomic Advisers' November 2013 forecasts and house price futures), around half of the gap could be closed by 2015 via a rise in the equilibrium debt ratio (Figure 4).[11]

Figure 4: Closing the debt gap from 2012Q4 to 2015Q4
(in percentage points)



Source: Authors' calculations.
Notes: The scenarios include actual data up to 2013Q3, and forecasts afterwards. The unemployment rate declines from 7.8% in 2012Q4 to 6.2% in 2015Q4, the house price-to-income ratio rises gradually by 8.2 percentage points between 2012Q4 and 2015Q4 (based on house price futures and MacroAdvisers' November 2013 forecasts) and, in Scenario 2, mortgage rates rise in line with the ten-year Treasury forward rates. The positive contribution from interest rates to the debt gap in the first scenario reflects the observed rise between 2012Q4 and 2013Q3. The homeownership rate and the share of the age group 35-54 in the total population remain constant over the horizon, while the loan-to-value ratio converges to its 1998-2012 average by the end of 2015.

[11] Under both scenarios, we assume that both the homeownership rate and the share of the 35-54 age group in the total population remain constant over the next years, while the loan-to-value ratio converges to its 1998-2012 average by the end of 2015.

On the other hand, the normalisation of monetary policy and exit from non-standard measures is likely to work in the opposite direction. Assuming a rise in mortgage rates in line with the ten-year Treasury yields as incorporated in the Treasury forward curve, a reversal of the low mortgage rates environment would act to push down, *ceteris paribus*, the level of sustainable debt, implying larger adjustment needs in actual debt. In this scenario, around one-third of the gap would be closed by 2015 via an increase in the equilibrium debt ratio.

4 Robustness of the results

The estimation results discussed in Section 3 should be interpreted with caution, particularly as the estimation of an unobservable variable (the equilibrium debt-to-income ratio) is surrounded by significant uncertainty. A number of potential econometric issues also poses challenges to the estimation: (1) the relationship between household debt and its long-term determinants could have changed over time; (2) the estimate of equilibrium debt conditional on house prices may be distorted by the housing boom; and (3) some explanatory variables may be endogenous. Moreover, there is more than one way to deal with cross-section dependence; in Section 4.4 we employ the Augmented Mean Group (AMG) estimator as an alternative to our baseline CCEPMG.

4.1 Model stability over time

Our first econometric concern is about the stability of the model over time. We acknowledge that, potentially, our baseline long-term coefficient estimates may be changing over time, reflecting structural changes in the relationship between household debt and its long-term determinants, possibly related to the Great Recession. In particular, we focus on house prices given its prominent role in driving the build-up of debt. We test for stability using two approaches.

In the first approach, we estimate recursively the baseline CCEPMG specification over different sample periods, similar to the technique applied in Arpaia and Turrini (2008). We start with the estimation over the 1999-2005 period, and then we add one year at a time until the entire sample period is covered. Figure C.2 in Appendix C plots the long-term coefficients of the baseline specification, together with the speed of adjustment and the Hausman test, over different samples. The results show that the long-term relationship between the household debt-to-income ratio and the six explanatory variables has changed over time. This is not surprising, as the financial crisis that erupted in early-2008 was unprecedented by the standards of previous recessions. Although demographics and the unemployment rate show a substantial time variation in their long-term relationship with debt, the other variables appear to be more stable. Moreover, even if there is some variation in the size of the coefficients of the latter variables, the coefficients' sign remains unchanged within the 95% confidence bands. The relationship between debt and house prices seems to have been temporarily distorted by the crisis period, with the coefficient rising when the sample starts to cover the financial crisis, but subsequently going back to the same coefficient as in the pre-crisis sample. More importantly, despite some evidence of the presence of structural changes in the aforementioned long-term relationship, the Hausman test performed recursively over the different sub-periods indicates that the CCEPMG assumption of long-run homogeneity is consistently not rejected.

In the second approach, we break down the sample into two non-overlapping sub-periods, estimating the model up to 2005Q4 – so as to exclude the crisis period – and from 2006Q1 until the end of 2012. Overall, this approach yields somewhat similar results as those from the recursive estimates. In particular, the long-term coefficients on house prices and interest rates are relatively stable over the two sub-periods, whereas the remaining ones are not (see columns (2) and (3) of Table 2). The estimated model in column (3) is qualitatively similar to the baseline one in column (1), even though it includes the crisis period. Finally, the Hausman test for both specifications again does not reject the assumption of long-run homogeneity of the coefficients across the US states. All in all, although there is evidence of some variation over time in the long-term relationship between debt and its determinants, probably as a consequence of the financial crisis of 2007-09, some coefficients appear to be rather stable, such as house prices and interest rates. In addition, the Hausman test suggests that it appears to be appropriate to assume a common long-run relationship across states.

Table 2: Alternative model specifications

	(1) Baseline CCEPMG	(2) Up to 2005Q4	(3) From 2006Q1	(4) Equil. HP	(5) AMG	(6) PMG	(7) MG	(8) 3 lags
Long-run coefficients								
House prices (HP)	24.320***	21.459***	23.949***	38.743**	40.988***	77.886***	79.660***	97.550***
	(4.818)	(3.781)	(4.588)	(18.195)	(4.203)	(3.391)	(8.300)	(3.702)
Homeownership rate	0.244***	0.318***	0.074	0.770***	0.026	0.476**	0.658**	-0.053
	(0.079)	(0.053)	(0.047)	(0.170)	(0.073)	(0.206)	(0.259)	(0.266)
Interest rates	-1.727***	-3.215***	-2.596***	-7.672***	-1.458***	-3.283***	-3.230***	-3.579***
	(0.370)	(0.329)	(0.411)	(1.125)	(0.427)	(0.813)	(1.060)	(0.906)
Unemployment rate	-1.277***	0.709**	-2.304***	-4.452***	-0.575***	-0.614*	-0.186	-3.725***
	(0.349)	(0.332)	(0.272)	(0.673)	(0.179)	(0.350)	(0.528)	(0.549)
35-54 age group	3.386***	-1.274	11.212***	-0.683	2.169***	-0.472	-1.489	-0.618
	(1.043)	(0.915)	(2.596)	(2.076)	(0.646)	(0.815)	(1.654)	(0.942)
Loan-to-value ratio	0.526***	-0.075	1.138***	0.755***	2.433***	0.683***	0.845***	0.093
	(0.087)	(0.058)	(0.065)	(0.231)	(0.168)	(0.217)	(0.242)	(0.281)
Speed of Adjustment	-0.378***	-0.909***	-0.616***	-0.155***	-0.803***	-0.144***	-0.292***	-0.136***
	(0.023)	(0.044)	(0.038)	(0.017)	(0.031)	(0.009)	(0.027)	(0.013)
Short-run coefficients								
Δ House prices	28.495***	2.036	33.880***	21.951***	4.549*	19.588***	11.292***	21.592***
	(5.132)	(8.046)	(6.055)	(5.639)	(2.749)	(2.582)	(2.779)	(3.134)
Constant	4.09	196.918***	-158.148***	-37.995***	-186.136***	-16.707***	-15.102	-6.172***
	(7.686)	(58.901)	(16.874)	(6.051)	(18.560)	(1.162)	(10.816)	(0.647)
Half-life	1.5	0.3	0.7	4.1	0.4	4.5	2	4.7
Observations	2,805	1,377	1,428	2,805	2,805	2,805	2,805	2,703
Hausman test	0.513	0.829	0.889	0.269	-	0.908	-	-

Notes: CCEPMG estimates corresponding to the common correlated effects specification of the PMG, with the exception of columns (6) to (8). The dependent variable is total household debt-to-income ratio. The lag structure (1 lag) was selected using the Schwartz Bayesian criterion. Foreclosures can only influence household debt in the short run. Due to space limitation, most of the short-run coefficients are omitted, with the exception of house prices. Standard errors are shown in parentheses. Asterisks, *, **, ***, denote, respectively, statistical significance at the 10, 5 and 1% levels. The half-life estimates indicate the number of quarters it takes to halve the gap between actual and equilibrium debt-to-income ratio. The Hausman test compares the PMG with the MG estimator and reports p-value under the null hypothesis that the PMG estimator is both efficient and consistent, i.e. that the long-run homogeneity restriction is valid. Model (2) was estimated on data up to 2005Q4, whereas model (3) takes into account the subsequent period (2006Q1 to 2012Q4). Model (4) uses a measure of equilibrium house price-to-income ratio. AMG stands for the Augmented Mean Group estimator. Estimate (8) refers to the standard PMG model with 3 lags, based on the Akaike Information Criterion.

4.2 The housing bubble

The second econometric issue relates to the estimated coefficients and resulting estimate of equilibrium debt conditional on house prices being potentially distorted by the recent speculative bubble in the US housing market. Although the issue at stake is similar in nature to the topic on the model stability over time, here we are interested mainly in adjusting the house price variable in order to better reflect a measure of "equilibrium" house prices, thus stripping out the "bubble" component.

It is a common practice to assess house price misalignments by comparing the evolution of house prices with respect to rents. To build our measure of equilibrium house prices, we take a path for house prices implied by a stable house price-to-rent ratio, in line with historical norms. It is computed only for the US aggregate, as rents are not available at the state level. The new CCEPMG estimate, displayed in column (4), shows that the results remain qualitatively similar compared with the baseline estimate. Although the coefficients on some variables are quantitatively somewhat different from the baseline estimate, one needs to be cautious in not over-interpreting this, as the significant rise in the standard errors shows that the precision of the estimates has declined. For instance, although the coefficient on house prices almost doubles from the baseline estimate, the standard error is 4 times larger. Interestingly, the fact that the speed of adjustment declines in magnitude suggests that, if, following a shock, house prices are not allowed to depart from its long-term average with respect to rents, the convergence of debt back towards its long-run equilibrium given by the explanatory variables is significantly slower.

4.3 Endogeneity

Finally, some of the regressors could be endogenous.[12] Pesaran and Shin (1999) show that, for inference on the long-run parameters, sufficient augmentation of the order of the ARDL model can correct for the problem of endogeneity of the explanatory variables. Hence, we augment our estimated model with more lags, in order to address the potential endogeneity issue. While the BIC suggests that one lag is sufficient (as used in the standard CCEPMG estimates reported in column (1) of Table 2), the AIC suggests an optimal lag order of 3 (see Table C.1 in Appendix C for the results). Column (8) therefore presents the PMG model augmented with three lags of the regressors and dependent variable to minimise the resultant endogeneity bias and to ensure that the regression residuals are serially uncorrelated.[13] The coefficients on house prices and the unemployment rate increase in magnitude, while the coefficients on the homeownership rate and the loan-to-value ratio lose their significance. Overall, however, the model remains rather robust to the inclusion of the additional lags.[14]

[12] While there are arguments to support the exogeneity assumption for some variables, such as demographics, the homeownership rate and the loan-to-value ratio – see Rajan (2010) for a discussion on changes in government policies and legislation to promote homeownership and enhance credit affordability, and Edelberg (2003) on the impact of technological progress in credit scoring techniques on credit availability – it could be the case that other variables, such as the house price-to-income ratio, are not free from endogeneity issues.

[13] We augment the PMG model, and not our baseline CCEPMG specification, with 3 lags, since adding 2 more lags to all the variables makes it problematic for the CCEPMG to be estimated. Therefore, the relevant comparison in this case is between the standard PMG and the PMG with 3 lags.

[14] Additionally, for the sake of completeness, in column (7) we also report estimates for the Mean Group (MG), which allows for heterogeneous long-run coefficients across US states, indicating that the estimated coefficients remain robust to this less restrictive assumption.

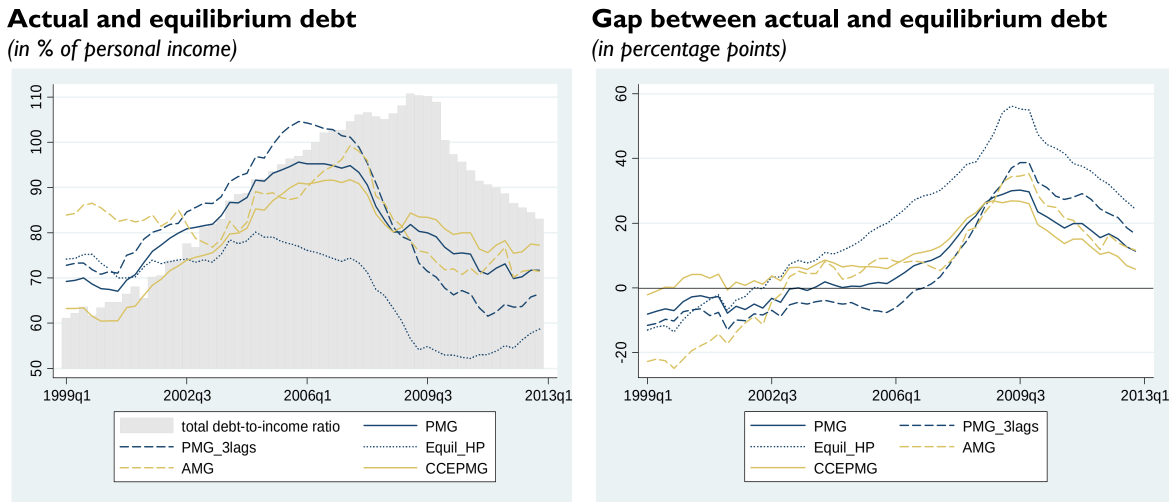
4.4 An alternative way to account for cross-section dependence

As mentioned before, there is more than one way to deal with cross-section dependence. Here, we employ the Augmented Mean Group (AMG) estimator, developed by Eberhardt and Teal (2010), as an alternative to our baseline CCEPMG. The AMG estimator is implemented in three steps: first, a pooled regression augmented with time dummies is estimated by first difference OLS. The coefficients on the (differenced) time dummies are estimated cross-group averages of the evolution of the unobserved effect over time (the so-called "common dynamic process"). The state-specific regression model is then augmented by this estimated process. Finally, the state-specific model parameters are averaged across states. As shown in column (5), the AMG model estimates yield practically the same results as those of the baseline CCEPMG.

4.5 Equilibrium debt and implied debt gaps for alternative specifications

In this section we present equilibrium debt and the implied debt gaps for a set of alternative specifications from Table 2, comparing them to our preferred CCEPMG model: the PMG model, the PMG augmented with 3 lags, the CCEPMG estimates using "equilibrium" house prices (assuming a stable house price-to-rent ratio, in line with its long-term average) and, finally, the AMG, an alternative estimator to the CCEPMG to correct for cross-section dependence across panel members. As mentioned previously, the resulting estimates need to be interpreted with some caution, given caveats involved in the estimation of an unobservable measure. Nevertheless, especially in qualitative terms, the conclusions reached are rather robust to alternative specifications, as shown in Figure 5.

Figure 5: Actual and equilibrium debt and implied gap for several specifications



Source: FRBNY/Equifax Consumer Credit Panel and authors' calculations.
Notes: Last observation refers to 2012Q4.

Table 3 provides a summary of our sensitivity analysis pointing to several conclusions which hold across the different estimation methods. Firstly, following the pronounced decline in the equilibrium debt ratio starting from around 2005-2007, more recently it has bottomed out and is on the rise again, with the trough estimated to have occurred between 2010Q4 and 2012Q1 depending on the different models used. Secondly, there has been a substantial widening in the debt gap since around 2005, which culminated in a peak (maximum deviation from the estimated equilibrium level) estimated to have occurred somewhere between 2008Q4 and 2009Q4; the peak ranged from 27 to 39 percentage points, according to four out of the five models. In the fifth specification – the model with equilibrium house prices – the estimated debt gap was generally larger throughout most of the sample period and the misalignment became more apparent at an earlier stage, since equilibrium debt started to fall already towards the end of 2004. This may be an indication that the rise in house prices up until 2006 might have helped to conceal in the other models the "true" misalignment of debt relative to its fundamentals. However, these results need to be interpreted with caution, since in this case the model has been estimated with a national, rather than a state, measure of equilibrium house prices due to data constraints. Finally, in the last several years, substantial progress has been achieved in the adjustment of household balance sheets. This is evidenced by the reduction in the estimated debt gaps to a range of 6-17 percentage points in 2012Q4, with, however, the model with equilibrium house prices standing at 24 percentage points. Based on all model results, more than half of the adjustment (ranging from 57% to 79%), relative to the respective peaks, has already been completed by that date.

Table 3: Estimates of equilibrium debt and debt gaps based on different specifications

MODEL SPECIFICATION	Equilibrium debt-to-income			Debt gap			Cumulative decline in the debt gap
	Recent trough		2012Q4	Peak		2012Q4	*% decline from peak to 2012Q4*
	pp	*timing*	*pp*	*pp*	*timing*	*pp*	
CCEPMG	75.4	2012q1	77.3	26.8	2008q4	5.7	78.6
CCEPMG equilibrium HP	52.2	2010q4	58.7	56.2	2009q2	24.3	56.7
AMG	70.1	2012q1	71.4	35.3	2009q4	11.6	67.0
PMG	69.8	2012q1	71.7	30.2	2009q3	11.4	62.4
PMG estimated with 3 lags	61.6	2011q2	66.4	38.7	2009q3	16.6	57.1

Source: Authors' calculations.

5 The heterogeneity of debt and deleveraging at the state level

The nature of the state-level data allows us to look beyond the national aggregate and provide further analysis of the heterogeneity of household deleveraging across US states. For instance, we can analyse the deleveraging developments in the "high deleveraging states", which include the 75th percentile of states with the largest declines in their household debt-to-income ratio from their respective peaks up to 2012Q4, and the "low deleveraging states", which comprise the 25th percentile of states with the smallest declines. The former group includes states like California, Arizona, Florida and Nevada, which have seen the sharpest housing and credit cycles. By contrast, the latter group is composed of states with more muted housing cycles, such as Kansas, Iowa or Arkansas. Overall, Table 4 indicates that the high deleveraging states, where households have undergone the largest amount of deleveraging, also had the highest debt ratios at their respective peaks. Meanwhile, the gap in 2012Q4 remained larger in the high deleveraging states, suggesting that, despite considerable debt reduction, the pressure on households to further repair their balance sheets remained more pressing in this group.

Additional regressions where the sample is split into high versus low deleveraging states are

Table 4: Household debt ratios and deleveraging indicators by state groups

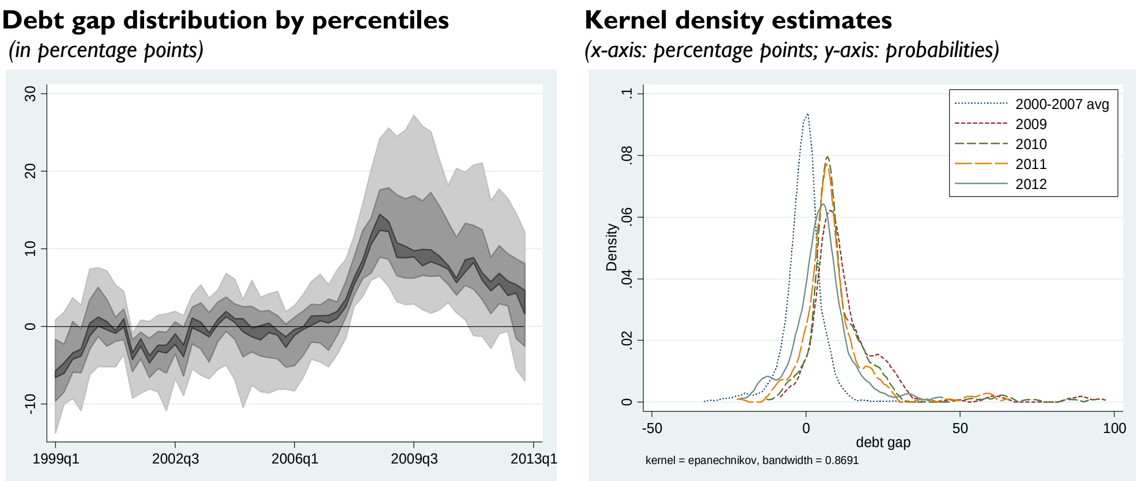
	Debt at peak	Debt in 2012Q4	Deleveraging from peak	Gap in 2012Q4
	% of income	*% of income*	*pp*	*pp*
		High deleveraging states		
75th percentile	129.2	92.7	-36.6	5.7
		Low deleveraging states		
25th percentile	75.7	67.4	-8.3	2.0

Source: FRBNY/Equifax Consumer Credit Panel and authors' calculations.

shown in Table C.6 of Appendix C. These suggest that the house price-to-income ratio has been an important driver in influencing debt-to-income in the long term in both sub-samples, albeit with a larger coefficient in the high deleveraging states (see columns (2) and (3)). This finding is confirmed by using the full sample, but interacting the house price variable with a dummy term that takes the value of 1 for high deleveraging states and zero otherwise (see column (4)). Furthermore, we have investigated whether the estimated models are different for "recourse" versus "non-recourse" states (see columns (5) and (6)). Non-recourse states refer to those states where the lender has no recourse against the borrower, if the borrower's house is sold at auction or via short sale for less than the amount owned by the lender. Overall, we find that the two models are similar to each other. In particular, the coefficient on house prices is very close to the one of the baseline CCEPMG.

The richness of the state-level data allows to look at the distribution of debt gaps across states, resulting from the CCEPMG model presented in Section 3.2. In particular, the share of states in need of further household deleveraging has been shrinking in the past three years (Figure 6).

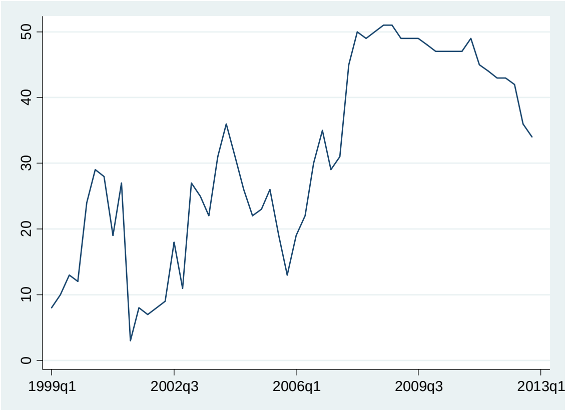
Figure 6: Debt gap distribution



Source: Authors' calculations.
Notes: The bands show the 10-90, 25-75 and 45-55 percentiles of the distribution. The chart on the right shows Kernel densities for the estimated debt gap at various points in time.

When the national debt gap was at its peak (2008Q4), virtually all US states had a deleveraging need in the household sector, meaning a debt gap above zero (Figure 7). The synchronised balance sheet adjustment carried out since then across all US states implied that, by 2012Q4, the debt gap for the states up to (and including) the 30th percentile was already below zero, suggesting that 1/3 of states (more precisely, 17 states) no longer faced deleveraging needs.

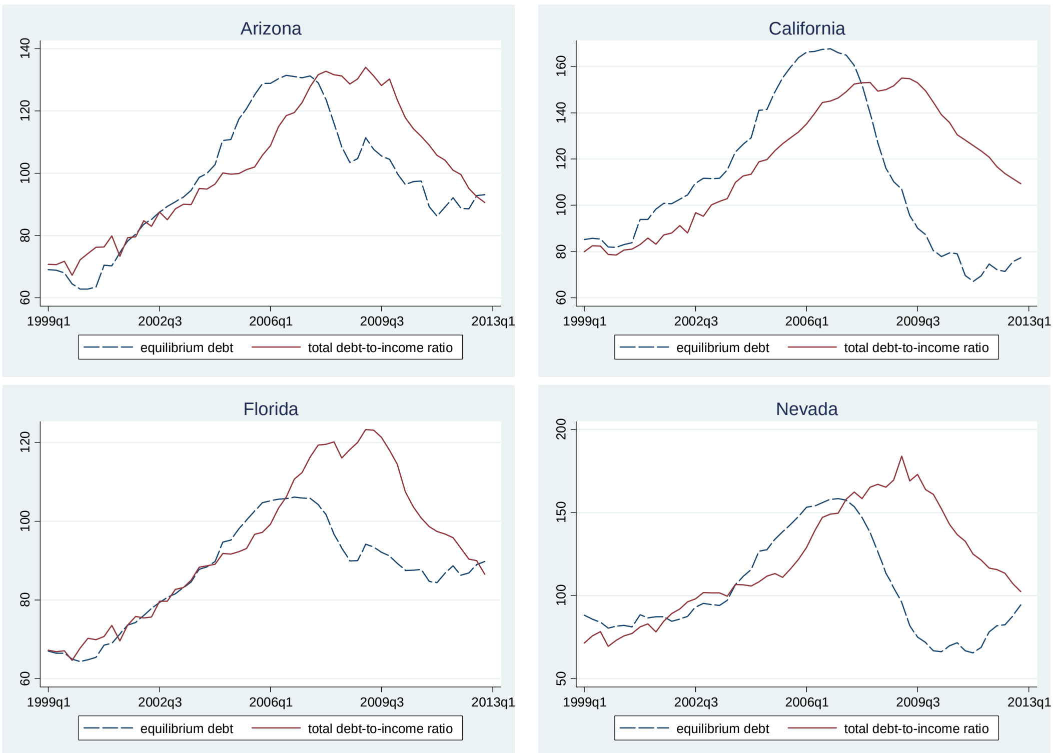
Figure 7: Number of states with debt gap above zero at each point in time



Source: Authors' calculations.

Those states that experienced pronounced boom-bust cycles in their housing markets (Arizona, California, Florida and Nevada) have managed to achieve substantial balance sheet repair within their household sectors (Figure 8). Nevertheless, for the states of California and, to a much lesser extent, Nevada, further adjustment still appeared to lie ahead.

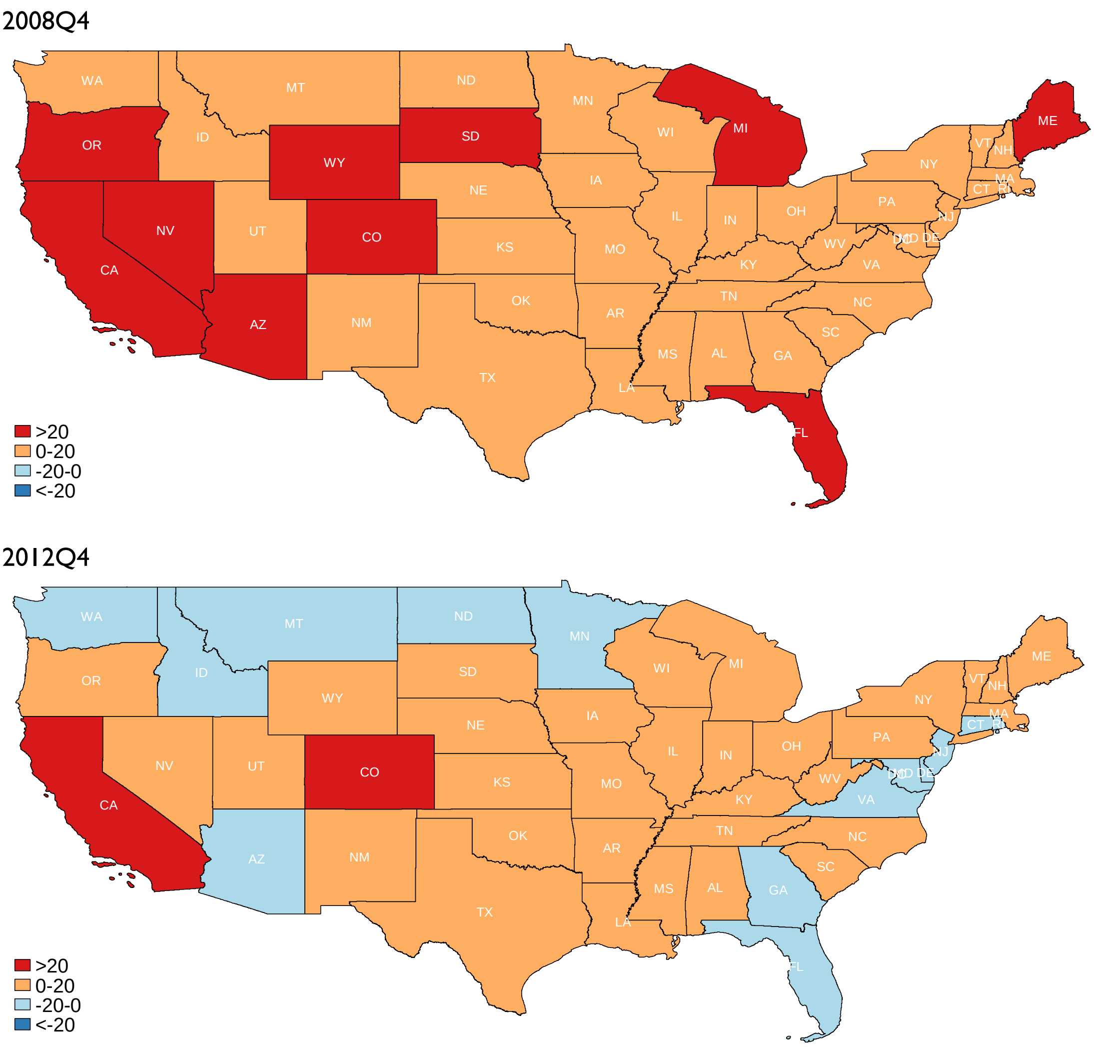
Figure 8: Actual and equilibrium debt in selected states



Source: Authors' calculations.
Note: Last observation refers to 2012Q4.

Figure 9 provides a spatial visualisation of estimated debt gaps across all US states (with the exception of Alaska and Hawaii) at two different points in time. First, in the final quarter of 2008, all US states faced either moderate or severe adjustment needs in the household sector. While for the majority of states, the actual debt-to-income ratio did not exceed the respective equilibrium level by more than 20 percentage points (states shown in orange colour), several other states were confronted with severe imbalances where the actual debt-to-income ratio exceeded by more than 20 percentage points the sustainable ratio (shown in red). By the end of 2012, the number of states with severe household debt imbalances has diminished markedly, while several states appeared no longer to face deleveraging pressures (shown in light blue).

Figure 9: Debt gaps across US states



Source: Authors' calculations.
Note: The debt gap measures the difference between the actual debt-to-income ratio and the estimated equilibrium ratio for each state based on the CCEPMG model. A positive debt gap indicates a need for balance sheet adjustment due to the actual ratio being above the estimated equilibrium ratio. The states of Alaska and Hawaii are not shown for convenience.

To cross-check our previous results at the state level, we adopt a related analysis in the spirit of Holly et al. (2010), who instead study the relationship between house prices and disposable income. As we have mentioned in Section 2.1, there is some evidence that the standard PMG specification may suffer from cross-section dependence, which motivated the use of CCEPMG, where the cross-section averages of the variables are acting as proxies for the unobserved common factors. Nevertheless, the factor loadings in Equation 2 of Section 2.1 differ across states, in the sense that the sensitivity of each state to the unobserved common factors is idiosyncratic. Following Holly et al. (2010), we can check how different these factor loadings are, by regressing the debt gap at the state level on the difference between average debt and the averages of the explanatory variables (taking into account the respective coefficients), which stands as a proxy for the unobserved common factor(s). More precisely, we estimate $(d_{it} - \theta_1' X_{it}) = \kappa_i + \chi_i(\overline{d}_t - \psi \overline{X}_t)$, where we are interested in the parameter χ_i.[15]

Table 5 presents the factor loadings by state, where a coefficient above 1 could be interpreted as an over-reaction of the debt gap of a particular state to the common unobserved factor(s). The factor loadings coefficients are rather heterogeneous, ranging from 0.26 to 2.33. There are two main findings from the table.

Table 5: Factor loadings estimates by state

State	Deleveraging from peak	Factor loadings	State	Deleveraging from peak	Factor loadings
Nevada	**-81.6**	**2.33* (0.11)**	New Mexico	-16.0	0.48 (0.03)
South Dakota	**-55.7**	**1.34* (0.10)**	New Jersey	-15.5	1.01 (0.03)
California	**-45.7**	**1.69* (0.06)**	**Massachusetts**	**-15.5**	**1.05* (0.02)**
Arizona	**-43.4**	**1.39* (0.05)**	**Rhode Island**	**-15.4**	**1.10* (0.04)**
Florida	**-36.8**	**1.28* (0.06)**	Tennessee	-15.3	0.73 (0.02)
Hawaii	-36.3	0.84 (0.10)	New York	-15.0	0.75 (0.02)
Oregon	**-28.1**	**1.24* (0.03)**	**Delaware**	**-14.8**	**1.36* (0.06)**
Maine	-26.4	0.93 (0.03)	Nebraska	-14.5	0.73 (0.02)
Vermont	-25.9	1.03 (0.04)	Texas	-13.5	0.45 (0.02)
Georgia	**-24.8**	**1.29* (0.03)**	Wisconsin	-13.1	0.75 (0.01)
Colorado	**-24.1**	**1.23* (0.06)**	Missouri	-12.4	0.72 (0.02)
Washington	-23.6	0.98 (0.03)	Pennsylvania	-12.3	0.63 (0.01)
Maryland	**-23.3**	**1.12* (0.02)**	Distr. of Columbia	-11.9	0.63 (0.05)
Michigan	**-23.0**	**1.13* (0.03)**	Alabama	-11.3	0.64 (0.03)
Virginia	**-22.5**	**1.16* (0.02)**	Kentucky	-10.8	0.78 (0.01)
New Hampshire	**-21.0**	**1.23* (0.04)**	Louisiana	-10.2	0.37 (0.03)
Wyoming	-19.8	0.50 (0.06)	Alaska	-10.1	0.56 (0.04)
Illinois	-19.5	1.02 (0.02)	Montana	-10.0	0.59 (0.03)
Idaho	-17.7	1.07 (0.04)	Oklahoma	-9.3	0.46 (0.02)
Ohio	-17.2	0.78 (0.04)	Kansas	-7.4	0.55 (0.02)
Minnesota	-17.1	0.99 (0.02)	Arkansas	-7.1	0.58 (0.02)
Utah	-17.1	0.80 (0.04)	North Dakota	-6.4	0.26 (0.01)
Indiana	-16.7	0.88 (0.03)	Mississippi	-6.3	0.56 (0.03)
North Carolina	-16.6	0.92 (0.02)	Iowa	-3.9	0.85 (0.02)
South Carolina	-16.5	0.70 (0.04)	West Virginia	-3.0	0.36 (0.04)
Connecticut	-16.0	0.87 (0.02)			

Notes: The figures on the column with the factor loadings refer to the slope coefficients obtained from regressing the debt gap of each state on the difference between average debt and the averages of the explanatory variables, which stands as a proxy for the unobserved common factor(s). More precisely, we estimate $(d_{it} - \theta_1' X_{it}) = \kappa_i + \chi_i(\overline{d}_t - \psi \overline{X}_t)$. Standard errors are shown in parentheses. The asterisk (*) denotes statistical significance at the 5% level for a test that the slope coefficient is above 1 (states shown in bold). Deleveraging from the peak takes into account the percentage point reduction in the debt-to-income ratio since the respective peak until 2012Q4.

First, US states that have reduced household debt more aggressively since the respective debt peak are also those that have higher coefficients. This suggests that states like Nevada, California, Arizona and Florida, which experienced boom-bust cycles in house prices, reacted more strongly to the unobserved common factor(s) at the US national level, which, in turn, might

[15] Based on Equation (3), $\psi = \frac{1}{N}\sum_{i=1}^{N} \frac{\beta_i}{-\alpha_i}$.

have been associated with the housing bubble and the financial crisis. The second finding is that we can broadly confirm our main results at the state level reported previously. In particular, states with a statistically significant coefficient above 1 (states shown in bold and with an asterisk) are also those where the actual debt-to-income ratio has deviated more markedly from the equilibrium level (as seen earlier in Figure 9), especially at a time when the debt gap at the national level was at its peak.

Finally, our dataset allows us to provide some insights into the heterogeneity of economic conditions across states, exemplified by differences in "high deleveraging" states (this time using a more restrictive definition, the 90th percentile in terms of household debt reduction undertaken) and "low deleveraging" states (10th percentile). Figure C.4 in Appendix C provides some support to the hypothesis that the overleveraging of households indeed has been an important determinant of economic performance. Not only did high deleveraging states experience much sharper boom-bust cycles in house prices and increases in delinquency rates, but they also had steeper declines in employment and stronger rises in their unemployment rates. Interestingly, mortgage interest rates have remained remarkably similar across all states, despite heterogeneous economic conditions. Exploring the link between the amplitude of the adjustment on the liability side of household balance sheets on the one hand, and economic activity on the other, remains a theme to be explored in future research.

6 Concluding remarks

The build-up in indebtedness in the US household sector since early-2000, and the subsequent balance sheet adjustment that began later in the decade, have been unprecedented by the standards of previous business cycles.

Our results show that the rise in actual debt at the national level resulted in a growing misalignment from the equilibrium level determined by fundamentals since around 2002-03. The trend was reinforced since mid-2007 by the decline in equilibrium debt, as fundamentals deteriorated. The deleveraging process, initiated in 2009, allowed the debt gap to be closed by roughly 80% from its peak until the end of 2012. Our baseline estimates indicate a remaining gap of around 6 percentage points by that time. At the state level, despite the synchronised balance sheet adjustment, deleveraging needs differ. We estimate that the debt adjustment appears to have been completed in one-third of the states by the end of 2012. The conclusions reached from our baseline model are rather robust to a set of alternative robustness checks.

While the additional balance sheet adjustment may take place via a further decrease in household debt relative to income, the time-varying nature of the equilibrium debt-to-income ratio in our models implies that the gap could also be closed by a rise in the estimated equilibrium. An improvement in economic conditions, such as a decline in the unemployment rate or rising house prices, would cause the sustainable debt ratio to rise, implying smaller adjustment needs. On the other hand, a normalisation of the monetary policy stance and a return to a higher interest rate environment would pose challenges to the deleveraging process in the future by reducing the level of sustainable household debt.

Our paper provides some interesting avenues for further work, namely by exploring the link between the amplitude of the adjustment on the liability side of household balance sheets on the one hand, and economic activity on the other. As we have shown in the last section of the paper, it seems that the extent of leveraging by households has indeed been correlated with economic performance across US states. In particular, high deleveraging states have experienced much sharper boom-bust cycles in house prices and increases in delinquency rates, while the labour market also had a worse performance. This topic remains a theme to be explored in future research.

Appendix

A Deriving the Common Correlated Effects Pooled Mean Group (CCEPMG) equation

We estimate a dynamic panel error correction model based on quarterly data for 50 US states (plus the District of Columbia) over the period 1999Q1-2012Q4, with subindices t = 1, 2, ..., T (T = 56) for quarters and i = 1, 2, ..., N (N = 51) for states. Following Pesaran et al. (1999), we assume an autoregressive distributed lag (ARDL) (1,1,1,...,1) dynamic panel specification of the form:

$$d_{it} = \mu_i + \rho_{i1} d_{i,t-1} + \delta'_{i0} X_{it} + \delta'_{i1} X_{i,t-1} + u_{it} \tag{1}$$

$$\text{where } X_{it} = \begin{bmatrix} hp_{it} \\ hown_{it} \\ i_{it} \\ ur_{it} \\ dem_{it} \\ ltv_{it} \\ frcl_{it} \end{bmatrix}, \ \delta_{i0} = \begin{bmatrix} \delta_{i01} \\ . \\ . \\ . \\ . \\ . \\ \delta_{i07} \end{bmatrix} \text{ and } \delta_{i1} = \begin{bmatrix} \delta_{i11} \\ . \\ . \\ . \\ . \\ . \\ \delta_{i17} \end{bmatrix}$$

If the variables are I(1) and cointegrated, then the error term is I(0) for all i.

Using $d_{it} = d_{i,t-1} + \Delta d_{it}$, (1) can be written as:

$$\Delta d_{it} = \mu_i + (\rho_{i1} - 1) d_{i,t-1} + \delta'_{i0} X_{it} + \delta'_{i1} X_{i,t-1} + u_{it} \tag{2}$$

Furthermore, since $X_{i,t-1} = X_{it} - \Delta X_{it}$, we can write the above equation as:

$$\Delta d_{it} = \mu_i + (\rho_{i1} - 1) d_{i,t-1} + (\delta'_{i0} + \delta'_{i1}) X_{it} - \delta'_{i1} \Delta X_{it} + u_{it} \tag{3}$$

To highlight the long-run relationship, we can write (3) in error-correction form:

$$\Delta d_{it} = \mu_i - (1 - \rho_{i1})[d_{i,t-1} - \tfrac{\delta'_{i0} + \delta'_{i1}}{1-\rho_{i1}} X_{it}] - \delta'_{i1} \Delta X_{it} + u_{it}$$

$$\Delta d_{it} = \mu_i + \phi_i (d_{i,t-1} - \theta'_{i1} X_{it}) + \delta^{*\prime}_{i1} \Delta X_{it} + u_{it} \tag{4}$$

$$\text{where } \phi_i = -(1 - \rho_{i1}),\ \theta_{i1} = \begin{bmatrix} \theta_{i11} \\ . \\ . \\ . \\ . \\ . \\ \theta_{i17} \end{bmatrix} = \begin{bmatrix} \frac{\delta_{i01} + \delta_{i11}}{1-\rho_{i1}} \\ . \\ . \\ . \\ . \\ . \\ \frac{\delta_{i07} + \delta_{i17}}{1-\rho_{i1}} \end{bmatrix} = -(\tfrac{\delta_{i0} + \delta_{i1}}{\phi_i}) \text{ and } \delta^*_{i1} = -\delta_{i1}$$

To estimate the parameters of the model, we apply the Pooled Mean Group (PMG) estimator as described in Pesaran et al. (1999). Under the PMG estimator assumption of long-run homogeneity, the long-run coefficients are assumed to be the same across states, i.e. $\theta_{i1} = \theta_1$. By contrast, the short-run coefficients and the group-specific error correction coefficients (the speed of adjustment) are allowed to differ across states, so that $\delta_{ij} = \delta_{ij}$, and $\phi_i = \phi_i$. In this case, the reported coefficient values are given by the means of the respective estimates for individual states. The PMG also assumes that the disturbances u_{it} are independently distributed across states i and time t with zero mean and state-specific variances $var(u_{it}) = \sigma_i^2$.

With the PMG assumption of long-run homogeneity, (4) can be written as:

$$\Delta d_{it} = \mu_i + \phi_i(d_{i,t-1} - \theta_1' X_{it}) + \delta_{i1}^{*\prime} \Delta X_{it} + u_{it} \tag{5}$$

$$\text{where } \theta_1 = \begin{bmatrix} \theta_{11} \\ . \\ . \\ . \\ . \\ . \\ \theta_{17} \end{bmatrix}$$

We assume that the foreclosure rate affects debt only in the short run but not in the long run (i.e. $\theta_{17} = 0$), so (5) can be written in its extended form as:

$$\begin{aligned} \Delta d_{it} = \mu_i + \phi_i(d_{i,t-1} - \theta_1 hp_{it} - \theta_2 hown_{it} - \theta_3 i_{it} - \theta_4 ur_{it} - \theta_5 dem_{it} - \theta_6 ltv_{it}) \\ + \delta_{i11}^* \Delta hp_{it} + \delta_{i12}^* \Delta hown_{it} + \delta_{i13}^* \Delta i_{it} + \delta_{i14}^* \Delta ur_{it} \\ + \delta_{i15}^* \Delta dem_{it} + \delta_{i16}^* \Delta ltv_{it} + \delta_{i17}^* \Delta frcl_{it} + u_{it} \end{aligned} \tag{6}$$

The term in brackets is the long-run relationship between debt and the explanatory variables, while $\theta_1, ..., \theta_6$ are the long-run coefficients, which are typically the object of primary interest (see Blackburne III and Frank 2007). ϕ_i is the speed of adjustment, which shows what percentage of the gap is being closed in each period and is expected to be negative and significant, if the variables are cointegrated and exhibit a return to long-run equilibrium. Furthermore, the average half-life of the adjustment for each state can be calculated using the formula $\log(0.5)/\log(1- \mid \phi_i \mid)$, see Durdu et al. (2013).

In the long-run, debt at the state level will be determined by:

$$d_{it} = \theta_{io} + \theta_{11} hp_{it} + \theta_{12} hown_{it} + \theta_{13} i_{it} + \theta_{14} ur_{it} + \theta_{15} dem_{it} + \theta_{16} ltv_{it} = \theta_{i0} + \theta_1' X_{it} \tag{7}$$

where $\theta_{i0} = \frac{\mu_i}{1-\rho_{i1}} = \frac{\mu_i}{-\phi_i}$ is the state-specific long-run constant.

As an alternative to our PMG estimator, we also apply the Mean Group (MG) estimator, which estimates independent error-correction equations for each state without imposing homogeneity restrictions on long-run effects, but rather computes the mean of estimated state-specific long-run coefficients. If the assumption of long-run homogeneity holds, the PMG estimates are both consistent and efficient, as "pooling" allows to sharpen the estimates compared to the MG. However, if the true model is heterogeneous, the PMG estimates are inconsistent, while the MG estimates are consistent in either case. We provide formal statistical evidence for choosing whether our PMG estimator is preferred to the MG estimator by applying a Hausman test on the homogeneity restriction that the long-run coefficient is the same for all states (see Pesaran et al. 1999).

To compute debt in the long run at the aggregate level, we apply the estimated long-run coefficients from the PMG, assumed to be homogeneous, to the US aggregate data:

$$d_t^{agg} = \overline{\theta}_0 + \theta_{11} hp_t^{agg} + \theta_{12} hown_t^{agg} + \theta_{13} i_t^{agg} + \theta_{14} ur_t^{agg} + \theta_{15} dem_t^{agg} + \theta_{16} ltv_t^{agg} = \overline{\theta}_0 + \theta_1' X_t^{agg} \quad (8)$$

The long-run aggregate constant term $\overline{\theta}_0$ is computed using the state-specific constants and the autoregressive coefficients on the lagged dependent variable as follows:

$$\overline{\theta}_0 = \frac{\sum_{i=1}^{N} \mu_i \rho_{i1}^0}{N} + \frac{\sum_{i=1}^{N} \mu_i \rho_{i1}^1}{N} + \frac{\sum_{i=1}^{N} \mu_i \rho_{i1}^2}{N} + \ldots + \frac{\sum_{i=1}^{N} \mu_i \rho_{i1}^\infty}{N} \quad (9)$$

which makes use of $\frac{\mu}{-\phi} = \mu \frac{1}{1-\rho_1} = \sum_{k=0}^{\infty} \mu \rho_1^k \iff |\rho| < 1$

The reason why $\overline{\theta}_0$ is not derived from the simple averages across states is because we cannot assume that the constant terms and the coefficients on the lagged dependent variables are independently distributed across panel members.

As discussed in the main text, the econometric estimation of Equation (5) may be affected by the presence of cross-section dependence across panel members. In what follows, we relax the assumption of cross-section independence of the error term of the standard PMG model by expressing the error term u_{it} as:

$$u_{it} = \lambda_i' f_t + \varepsilon_{it} \quad (10)$$

where an unspecified number of unobserved common factors f_t with idiosyncratic factor loadings λ_i are allowed to capture time-variant heterogeneity and cross-section dependence, while ε_{it} are now idiosyncratic errors independently distributed across i and t. In this set-up, the factors f_t can be non-linear and also non-stationary. The regressors X_{it} are allowed to be driven by some of the same common factors as the dependent variable.

We employ two alternative estimators that have been developed to allow for correlation across panel members due to unobserved common time-specific effects as described above: the Augmented Mean Group (AMG) estimator introduced in Eberhardt and Teal (2010) and the Common Correlated Effects Pooled Mean Group (CCEPMG) estimator (see Pesaran 2006, Binder and Offermanns 2007, Chudik and Pesaran 2013).

The basic procedure behind the AMG estimator by Eberhardt and Teal (2010) is described in the main text and is applied to estimate Equation (3), i.e. prior to imposing the long-run homogeneity restriction ($\theta_{i1} = \theta_1$). The reported long-run coefficients in Table 2 of the main text are averaged across panel members (as usual under the Mean Group type estimators). Their standard errors have been computed with the use of the delta method. These long-run coefficients are then employed to compute equilibrium debt and the associated debt gaps at the national level for the AMG specification.

The other method used to correct for cross-section dependence in the disturbances is the CCEPMG estimator, which is chosen as our preferred specification when reporting the main results. It is based on the Common Correlated Effects Mean Group (CCEMG) estimator developed by Pesaran (2006). The basic idea is to filter the individual-specific regressors in a way that the differential effects of unobserved common factors are eliminated. The CCEMG solves the problem by augmenting the regressors in the group-specific regression equation with cross-section averages of the dependent variable and the individual-specific regressors.

The focus of the CCEMG estimator is on obtaining consistent estimates of the parameters related to the observable variables, while the estimated coefficients on the cross-section averaged variables are not interpretable in a meaningful way: they are merely present to filter out the biasing impact of the unobservable common factor (see Eberhardt 2012). The CCEMG estimator is robust to the presence of a limited number of "strong" factors (which can represent global shocks, possibly related to the recent financial crisis or the housing bubble), as well as an infinite number of "weak" factors possibly associated with local spillover effects. Moreover, the CCEMG estimator is robust to nonstationary common factors (Kapetanios et al. 2011) and it continues to hold under slope homogeneity and in the presence of any fixed number of unobserved factors, which is important in practical applications (see Pesaran 2006).

The common correlated effects (CCE) augmentation has been applied empirically in conjunction with the PMG estimator in Binder and Offermanns (2007). The CCEMG estimator has been further extended in Chudik and Pesaran (2013) to allow for the inclusion of lagged values of the dependent variable among the regressors. In the latter case, it has been shown that the CCEMG continues to be valid as long as: (1) a sufficient number of lags of cross-section averages are included in individual equations of the panel, and (2) the number of cross-section averages is at least as large as the number of unobserved common factors minus one (m-1). Chudik and Pesaran (2013) point out that in practice, where the number of unobserved common factors is unknown, it is sufficient to assume that the number of available cross-section averages is at least $m_{\max} - 1$, where $m_{\max}$ denotes the assumed maximum number of unobserved factors. Based on the existing literature, in most empirical applications $m_{\max}$ is expected to be small, as only few factors (most likely no more than two or three) tend to explain much of the predictable variations (see Stock and Watson 2002, Giannone et al. 2005, Bai and Ng 2007).

Augmenting the standard PMG model in Equation (5) with cross-section averages as discussed above, our equation for the CCEPMG estimator becomes:

$$\Delta d_{it} = \mu_i + \phi_i(d_{i,t-1} - \theta_1' X_{it}) + \delta_{i1}^{*\prime} \Delta X_{it} + \alpha_i \overline{d}_t + \beta_i \overline{X}_t + \gamma_i \Delta \overline{d}_t + \eta_i \Delta \overline{X}_t + \varepsilon_{it} \qquad (11)$$

where $\overline{d}_t$ and $\overline{X}_t$ are averages of the dependent variable and the regressors across states, computed at every time period t.

Following (11), the long-run relationship between d and X at the state level is given by:

$$d_{it} = \theta_{i0} + \theta_1' X_{it} + \vartheta_i \overline{d}_t + \psi_i' \overline{X}_t \qquad (12)$$

$$\text{where } \vartheta_i = \frac{\alpha_i}{-\phi_i} \text{ and } \psi_i = \begin{bmatrix} \psi_{i1} \\ . \\ . \\ . \\ . \\ . \\ \psi_{i7} \end{bmatrix} = \begin{bmatrix} \frac{\beta_{i1}}{-\phi_i} \\ . \\ . \\ . \\ . \\ . \\ \frac{\beta_{i7}}{-\phi_i} \end{bmatrix} = -(\frac{\beta_i}{\phi_i})$$

In our final CCEPMG specification, the CCE augmentation of the standard PMG model employs the following cross-section averages (included both in levels and in differences): log of house price to income ratio (*hp*), unemployment rate (*ur*), and 35-54 age group (*dem*). The choice to use a subset of the available variables for the augmentation is driven by the need to keep the model relatively parsimonious, as well as by concerns about possible misspecification resulting from multicollinearity issues in the case that cross-section averages for all the available variables are used, given that some are either available only at the national level (the loan-to-value ratio, *ltv*), or display little variability at the cross-section dimension (mortgage interest rates, *i*).

Based on our final CCEPMG specification, long-run debt at the US aggregate level will be given by:

$$\begin{aligned} d_t^{agg} &= \overline{\theta}_0 + \theta_{11} hp_t^{agg} + \theta_{12} hown_t^{agg} + \theta_{13} i_t^{agg} + \theta_{14} ur_t^{agg} + \theta_{15} dem_t^{agg} \\ &+ \theta_{16} ltv_t^{agg} + \sum_{j=0}^{\infty} \psi_j \overline{X}_{t-j} = \overline{\theta}_0 + \theta_1' X_t^{agg} + \sum_{j=0}^{\infty} \psi_j \overline{X}_{t-j} \end{aligned} \tag{13}$$

where $\overline{\theta}_0$ is computed as described in (9), while the contribution from the cross-section averages $\sum_{j=0}^{\infty} \psi_j \overline{X}_{t-j}$ uses $\psi_j = \frac{\sum_{i=1}^{N} \beta_i \rho_{i1}^j}{N}$ based on the approach used for the aggregation of the long-run constant term.

For practical reasons, in the main text we report the estimates for equilibrium debt and debt gaps using the term $\sum_{j=0}^{\infty} \psi_j \overline{X}_t$(without lagged values of the cross-section averages), so as to have estimated values for the whole sample period. We do this after having checked that the results are rather similar to the ones using the term $\sum_{j=0}^{\infty} \psi_j \overline{X}_{t-j}$, computed with up to 30 lags for the term $\overline{X}_{t-j}$ (results are available upon request).

B Data sources and descriptive statistics

Variables used (state-level or national-level in parenthesis):

- **d - debt-to-income ratio (state)**: aggregated state-level household debt data from the FRB of New York/Equifax (quarterly) divided by personal income from the Bureau of Economic Analysis (quarterly).

- **hp - log of house price-to-income ratio (state)**: FHFA house price index from the Federal Housing Finance Agency (quarterly) divided by personal income per capita, in order to avoid distortions due to state size (quarterly). Population is taken from the Bureau of Economic Analysis (quarterly interpolated from annual data).

- **hown - homeownership rate (state)**: owner-occupied housing units divided by total occupied housing units from the Bureau of the Census (quarterly).
- **i - interest rates on conventional mortgages (state)**: effective interest rate on conventional home mortgages from the Federal Housing Finance Board (quarterly interpolated from annual data).
- **ur - unemployment rate (state)**: unemployment rate from the Bureau of Labor Statistics (quarterly).
- **dem - demographics: share of 35-54 age group to total population (state)**: resident population by age group from the Bureau of the Census (quarterly interpolated from annual data).
- **ltv - loan-to-value ratio (national)**: loan-to-value ratio on conventional mortgages for previously occupied homes (excluding refinancing loans) from the Federal Housing Finance Agency (quarterly).
- **frcl - foreclosure rate, inventory (state)**: total number of loans in the legal process of foreclosure as a percentage of the total number of mortgages in the pool during a quarter, taken from the Mortgage Bankers Association (quarterly).

Table B.1: Descriptive statistics

Variable	Obs.	Mean	Std. Dev.	Min	Max
Total debt to-income ratio	2856	81.3	22.2	37.7	184.1
Log of house price-to-income ratio	2856	1.9	0.2	1.3	2.6
Homeownership rate	2856	69.3	6.2	37.6	82.4
Interest rates	2856	6.1	1.0	3.6	8.5
Unemployment rate	2856	5.7	2.1	2.2	14.1
35-54 age group	2856	28.7	1.7	23.0	33.3
Loan-to-value ratio	2856	76.0	2.0	72.1	80.4
Foreclosure rate	2856	1.9	1.6	0.2	14.5

Source: Bureau of Economic Analysis, Bureau of Labor Statistics, Census Bureau, Federal Housing Finance Agency, Federal Housing Finance Board, FRBNY/Equifax Consumer Credit Panel, Mortgage Bankers Association, and authors' calculations.

C Additional tables and figures

Table C.1: Lag order selection based on the Bayesian and Akaike information criteria

Number of states for which the respective lag order is chosen	Lag order 1	2	3
BIC	26	10	15
AIC	1	4	46

Note: BIC and AIC stand respectively for Bayesian and Akaike information criteria.

Table C.2: Unit-root tests for variables available at the state level (p-values)

		Debt-to-income ratio	House price-to-income ratio	Homeowner. rate	Interest rate	Unempl. rate	35-54 age group
Levin-Lin-Chu	No constant	1.000	0.001	0.226	0.000	0.004	0.000
	With constant	0.000	1.000	0.000	1.000	0.000	1.000
	No means	0.215	0.001	0.000	0.000	0.000	0.026
Breitung	No constant	1.000	0.001	0.228	0.000	1.000	0.000
	With constant	0.962	0.416	0.000	1.000	0.867	1.000
	No means	0.001	0.998	0.000	0.004	0.672	1.000
	Robust	0.554	0.445	0.000	0.998	0.389	1.000
Im-Pesaran-Shin	Uncorr. errors	0.000	1.000	0.000	1.000	1.000	1.000
	No means	0.067	0.408	0.000	1.000	0.999	1.000
	Correl. errors	0.574	0.088	0.000	0.000	0.273	0.995
Fisher	ADF	0.001	1.000	0.066	1.000	0.129	1.000
	PP	0.000	1.000	0.000	1.000	1.000	1.000
	ADF (no means)	0.865	0.133	0.000	0.000	0.624	0.997
	PP (no means)	0.254	0.422	0.000	0.463	0.808	1.000
I(1) at the 1% level		64%	79%	21%	57%	79%	86%

Notes: The tests are based on the null hypothesis that the variables are I(1).

Table C.3: Unit-root tests for variables available at the national level (loan-to-value ratio)

		T-statistic	Critical value at 1%
Augmented Dickey-Fuller	No constant	-0.413	-2.616
	With constant	-1.952	-3.567
	With drift	-1.952	-2.394
	With 3 lags	-2.286	-3.572
Phillips-Perron	No constant	-0.391	-2.616
	With constant	-2.145	-3.567
	With trend	-2.132	-4.130
Kwiatkowski-Phillips-Schmidt-Shin	Trend stationarity	0.109	0.216
	Level stationarity	0.172	0.739
I(1) at the 1% level		78%	

Notes: The first two tests are based on the null hypothesis that the variable contains a unit root. By contrast, the KPSS test is based on a null hypothesis of stationarity.

Table C.4: Panel cointegration test (τ-bar statistic)

		With constant	With constant and trend
Lag order	1	-5.237***	-5.346***
	2	-3.978***	-4.077*
	3	-3.176	-3.297
	4	-2.039	-4.190**

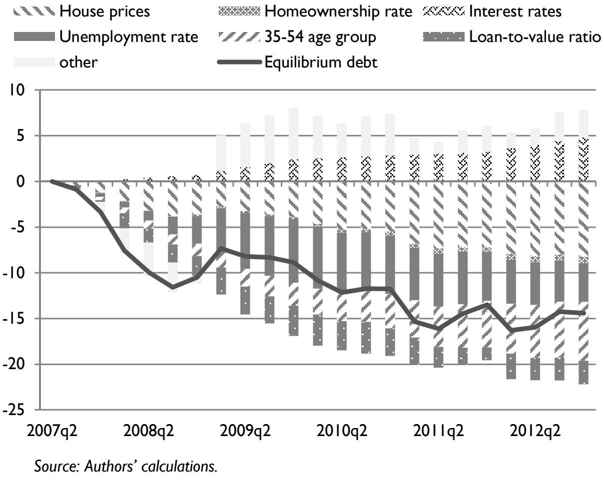
Notes: The test is based on Gengenbach et al. (2009), who have developed a second-generation panel cointegration test that takes into account cross-section dependence. The results test the null hypothesis of no cointegration between the debt-to-income ratio and the five explanatory state-level variables in our error-correction model, with the exception of the loan-to-value ratio which is available only at the national level. The reported lag order is based on the model specification in levels. Asterisks *, ** and *** denote, respectively, significance at the 10%, 5% and 1% levels based on the critical values from Gengenbach et al. (2009), Table 3, p. 31.

Table C.5: Unit-root tests on the CCEPMG state-level debt gaps

		p-values
Levin-Lin-Chu	No constant	0.000
	With constant	0.000
	No means	0.388
Breitung	No constant	0.000
	With constant	0.080
	No means	0.006
	Robust	0.208
Im-Pesaran-Shin	Uncorr. errors	0.000
	No means	0.001
	Correl. errors	0.005
Fisher	ADF	0.004
	PP	0.000
	ADF (no means)	0.107
	PP (no means)	0.001
I(0) at the 1% level		71%

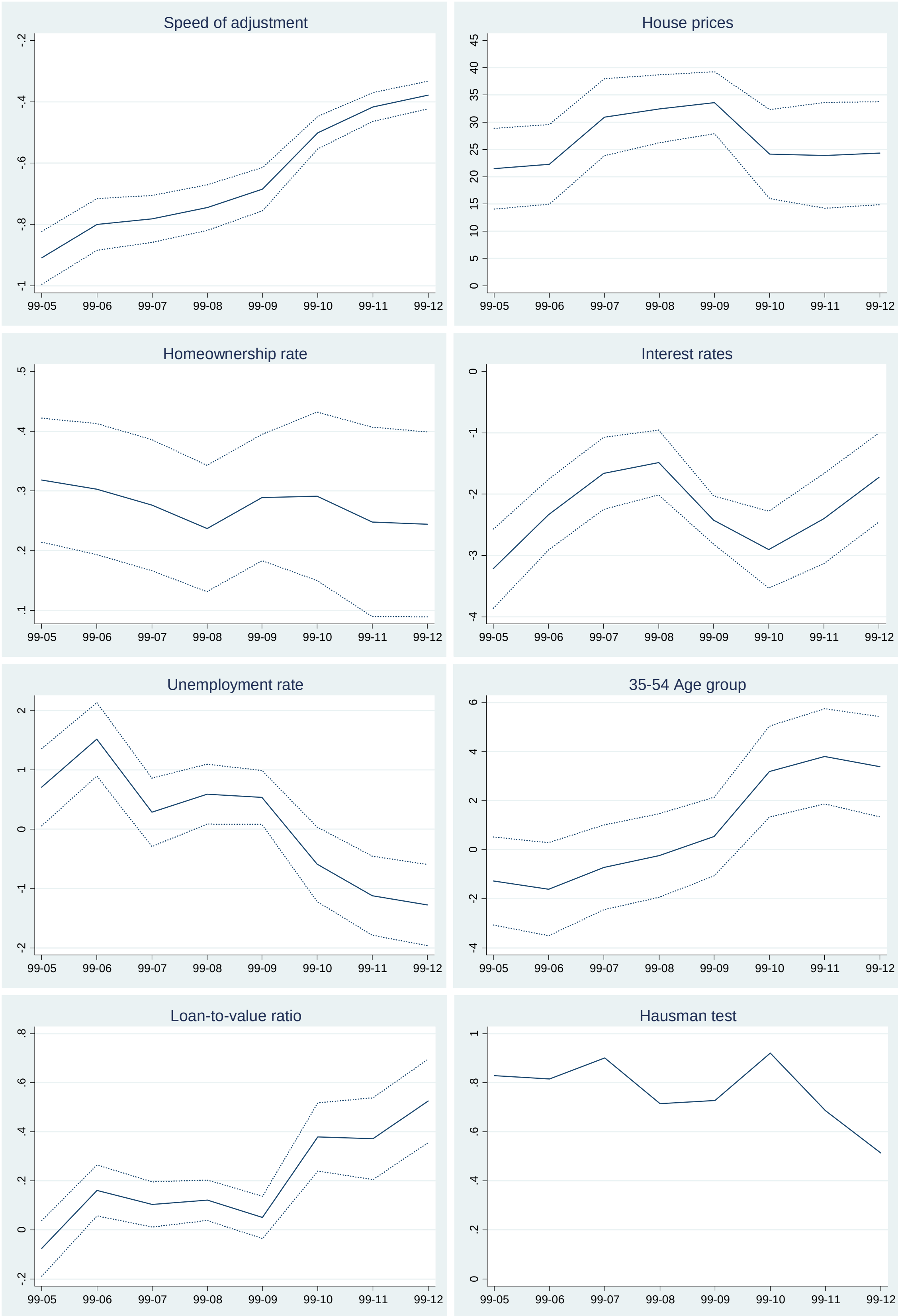
Notes: The tests are based on the null hypothesis that the variable contains a unit root.

Figure C.1: Decomposing the decline in CCEPMG equilibrium debt from 2007Q2



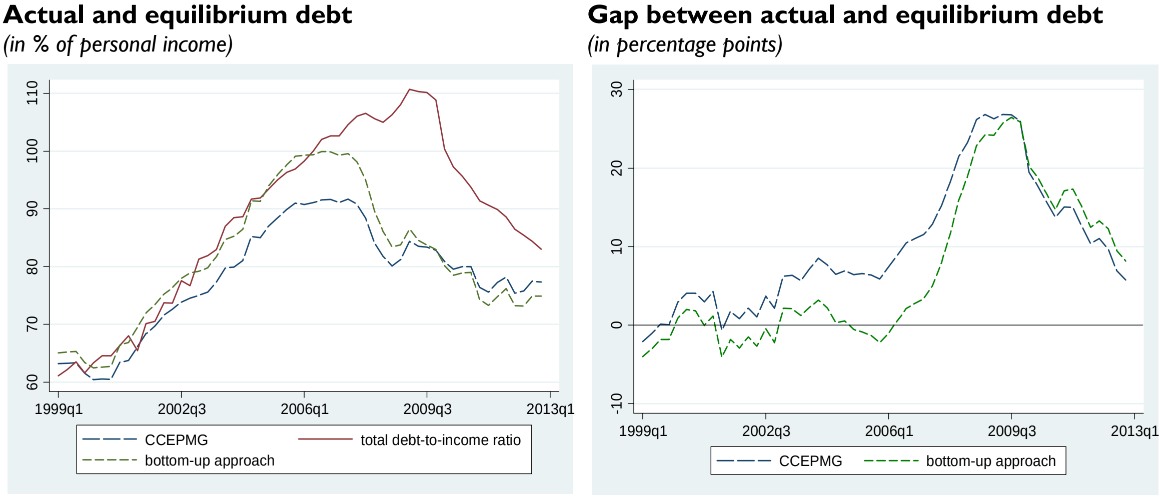
Source: Authors' calculations.

Figure C.2: Recursive CCEPMG estimates using different sample periods



Note: The CCEPMG is estimated recursively over different sample periods: the first estimate considers the 1999-2005 period, and in the subsequent estimates it adds one year at a time. The speed of adjustment corresponds to the error correction term in the CCEPMG model. The Hausman test reports p-value under the null hypothesis that the CCEPMG estimator is both efficient and consistent, i.e. that the long-run homogeneity restriction is valid. The remaining solid lines refer to the long-run coefficients on the variables included in the CCEPMG model. Bands around the point estimates consider +/- 2 standard errors.

Figure C.3: Equilibrium debt and implied gap using a bottom-up approach



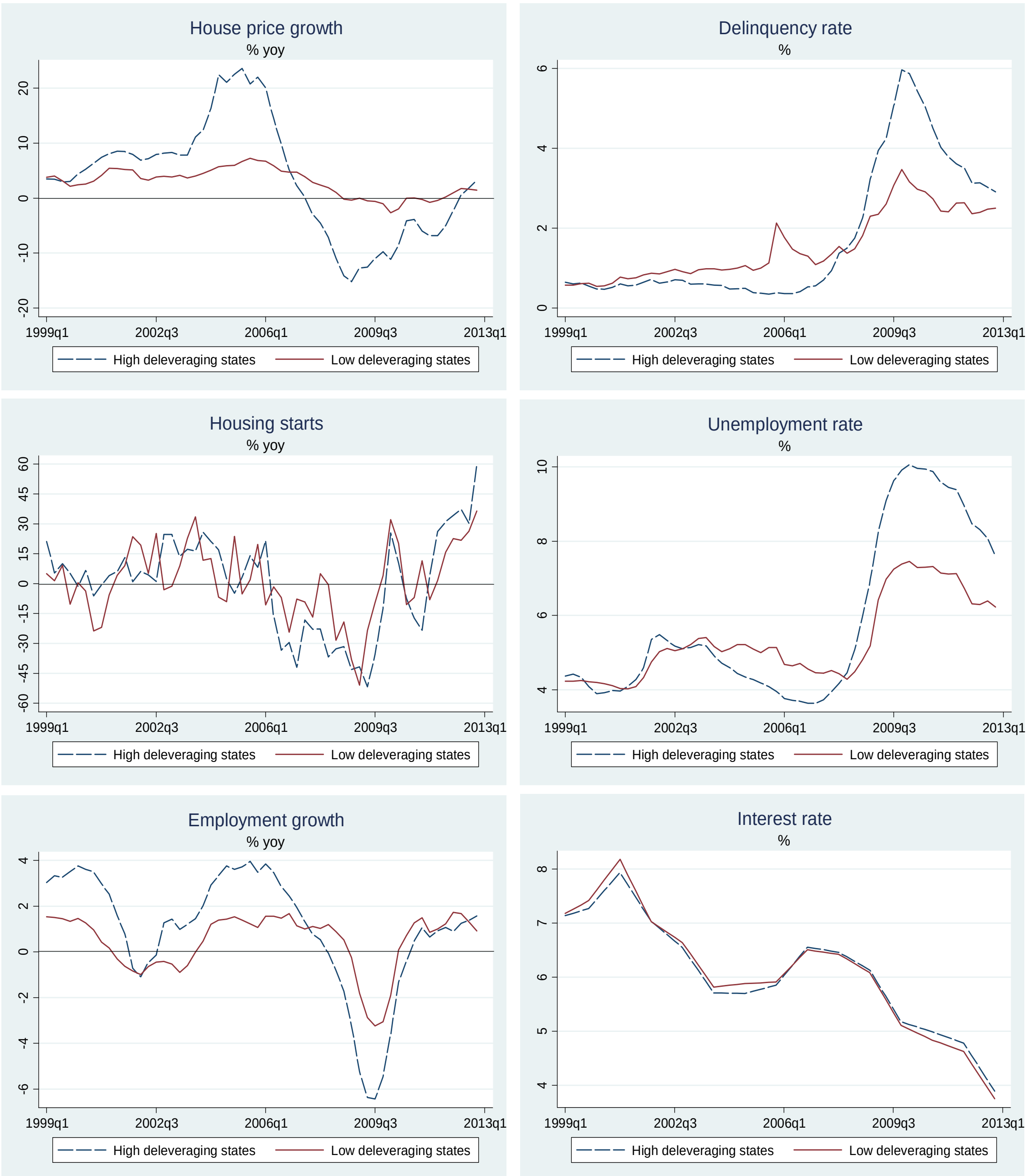
Source: FRBNY/Equifax Consumer Credit Panel and authors' calculations.
Notes: Equilibrium debt in the bottom-up approach is computed from the aggregation of state-level equilibrium estimates, derived by taking state-weights for income. Last observation refers to 2012Q4.

Table C.6: Sensitivity analysis of CCEPMG model with split regressions and interaction terms

	(1) CCEPMG	(2) HD states	(3) LD states	(4) House prices	(5) Non-Rec. states	(6) Recourse states
Long-run coefficients						
House prices (HP)	24.320***	59.806***	19.801***	20.417***	23.714***	20.478***
	(4.818)	(16.479)	(7.548)	(5.024)	(7.894)	(5.891)
Homeownership rate	0.244***	0.708**	0.115	0.245***	0.069	0.326***
	(0.079)	(0.336)	(0.095)	(0.079)	(0.121)	(0.094)
Interest rates	-1.727***	-0.635	-2.043***	-1.775***	-1.195**	-1.836***
	(0.370)	(1.221)	(0.516)	(0.369)	(0.609)	(0.440)
Unemployment rate	-1.277***	-1.085	-2.391***	-1.271***	-1.236*	-1.465***
	(0.349)	(1.374)	(0.468)	(0.348)	(0.751)	(0.379)
35-54 age group	3.386***	9.634**	3.518***	3.269***	-2.718	5.931***
	(1.043)	(4.529)	(1.290)	(1.045)	(1.969)	(1.222)
Loan-to-value ratio	0.526***	0.561*	0.583***	0.525***	0.457***	0.439***
	(0.087)	(0.298)	(0.123)	(0.087)	(0.145)	(0.103)
HP*HD states				46.539***		
				(16.249)		
Speed of Adjustment	-0.378***	-0.321***	-0.497***	-0.380***	-0.390***	-0.383***
	(0.023)	(0.048)	(0.048)	(0.023)	(0.062)	(0.026)
Half-life	1.5	1.8	1	1.4	1.4	1.4
Observations	2,805	715	715	2,805	660	2,145
Hausman test (p-value)	0.513	0.998	0.637	0.52	0.535	0.655

Notes: CCEPMG estimates where the dependent variable is household debt-to-income ratio. The lag structure (1 lag) was selected using the Schwartz Bayesian criterion. Foreclosures can only influence household debt in the short run. Standard errors are shown in parentheses. Asterisks, *, **, ***, denote, respectively, statistical significance at the 10, 5 and 1% levels. The half-life estimates indicate the number of quarters it takes to halve the gap between actual and equilibrium debt-to-income ratio. The Hausman test reports p-value under the null hypothesis that the CCEPMG estimator is both efficient and consistent, i.e. that the long-run homogeneity restriction is valid. "High deleveraging states" (HD) stands for the 75th percentile of states with the largest declines in their household debt-to-income ratio from their respective peaks up to 2012Q4, while the "low deleveraging states" (LD) are the 25th percentiles of states with the smallest declines. "Non-recourse states" refer to those states where the lender has no recourse against borrowers if the borrowers' house is sold at auction or via short sale for less than the amount owned by the lender (Alaska, Arizona, California, Connecticut, Idaho, Minnesota, North Carolina, North Dakota, Oregon, Texas, Utah and Washington, D.C.)

Figure C.4: Average developments in economic indicators for high vs low deleveraging states



Note: "High deleveraging states" are those states that featured the largest declines in their household debt-to-income ratios between the peak for each state and 2012Q4, defined by the 90th percentile. These include Arizona, California, Florida, Hawaii, Nevada and South Dakota. The "low deleveraging states" are those that featured the smallest declines, defined as the 10th percentile and include Arkansas, Iowa, Kansas, Mississippi, North Dakota and West Virginia.

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