Abstract — This paper investigates the impact of changing housing rental rates upon aggregate consumption based on yearly data for 18 OECD countries observed between 1970 and 2004. Estimates of long run elasticities are derived from cointegrating relationships with panel estimation techniques and compared with estimates implicitly given in short run dynamic equations. The results across all estimation approaches yield similar magnitudes and indicate a significantly negative impact of rental rate increases upon aggregate consumption. For the EU-25 or the USA, for example, relevant elasticities lie in the range between -0.13 and -0.17. The jointly estimated elasticities of the wealth components seem unaffected, relative to previous studies, by the inclusion of the rental housing market variables.

Keywords: housing rental rates, aggregate consumption, panel estimation, wealth effects

JEL-Classification: E21, R21, C33
The housing rental rate elasticity of aggregate consumption: A panel study for OECD countries

Abstract — This paper investigates the impact of changing housing rental rates upon aggregate consumption based on yearly data for 18 OECD countries observed between 1970 and 2004. Estimates of long run elasticities are derived from cointegrating relationships with panel estimation techniques and compared with estimates implicitly given in short run dynamic equations. The results across all estimation approaches yield similar magnitudes and indicate a significantly negative impact of rental rate increases upon aggregate consumption. For the EU-25 or the USA, for example, relevant elasticities lie in the range between -0.13 and -0.17. The jointly estimated elasticities of the wealth components seem unaffected, relative to previous studies, by the inclusion of the rental housing market variables.

Keywords: housing rental rates, aggregate consumption, panel estimation, wealth effects

JEL-Classification: E21, R21, C33
1 Introduction

This paper contains an empirical investigation of the impact of changing housing rental rates upon aggregate consumption. Roughly a third of all households in the EU-25 and the USA live in rented accommodation (see Catte et al. [2004b] or Ball [2005]) and typically pay 20-30% of their income for rent (e.g. Genesove [2003]). Therefore changing rental rates may be associated with significant economic effects. From this point of view it is surprising that the empirical macroeconomic literature on housing has so little to say about rental housing (see the last literature review of Leung [2004]).

As Whitehead [1999] states in her survey article, “housing markets are inherently imperfect” due to some distinct characteristics of housing. Along with other factor this “motivates a range of policy interventions” (HM Treasury [2005]) in housing markets around the globe. These interventions occur in the form of particular taxation laws, mortgage subsidization policies, rent controls, zoning laws and many others (see for example ECB [2003], Catte et al. [2004a] or OECD [2004]). For this reason prevailing rental rates, particularly relative to the price of owner occupied housing, are not just the outcome of the market process but at least as much the result of the influence of public policy towards housing. Hence an understanding of the macroeconomic consequences of changing rental rates is important for assessments of housing related policy measures.

The impact of changing rental rates upon aggregate consumption is a priori unclear, because tenants and landlords are affected differently: The former lose in terms of real income and are likely to lower overall consumption in real terms, the latter gain. Furthermore rental rates are linked to house prices and thus rental changes may affect landlords consumption also via the wealth channel. Second round effects via changing incomes of producers of goods other than rental housing services add further to this indeterminacy. Therefore the net effect upon aggregate consumption can only be determined empirically.

The renter share varies strongly between countries, being for example below 20% for Spain and above 50% for Germany today. Therefore one would expect the rental rate impact upon consumption to differ in strength across countries, even without other structural differences between countries. For the purpose of panel estimation this requires modeling
corresponding elasticities as country specific.

The remainder of the paper is organized as follows: Section 2 motivates the regression setup and discusses related econometric issues, section 3 describes the data used, section 4 presents the estimation results and section 5 concludes. Relevant literature is noted along the way.

2 Estimation

2.1 Regression setup

The starting point and main reference for the present paper is the literature on real estate wealth effects upon consumption, e.g. Boone & Girouard [2002], Dvornak & Kohler [2003], Ludwig & Sløk [2004], Catte et al. [2004b], Case et al. [2005], Al-Eyd et al. [2005] or Bostic et al. [2005]. The present paper deviates from this literature by adding two explanatory variables to the aggregate consumption function to reflect basic rental housing market conditions. As a first attempt to investigate the potential role of rental housing for aggregate consumption, the chosen specification is rather parsimonious\(^1\) and ignores other potentially relevant issues.\(^2\) The (shortrun) equation to be estimated is, therefore, simply specified as

\[
C_{it} = \alpha_1 C_{i,t-1} + \alpha_2 Y_{it} + \alpha_3 P^H_{it} + \alpha_4 P^S_{it} + \left( \alpha_5 S^R_{it} + \alpha_6 S^R_{it} \right) + \nu_i + \theta_t + \epsilon_{it},
\]

(1)

where \(C\) denotes consumption per capita, \(Y\) disposable income per capita, \(R\) the housing rental rate, \(P^H\) house prices and \(P^S\) stock prices. All these variables are real (i.e. deflated by a consumer price index) and in log terms. \(S \in [0,1]\) denotes the renter share and subscripts \(i\) and \(t\) indicate countries resp. time periods.

Apart from the framed expression formulation (1) corresponds to a standard model used

---

\(^1\) An extra benefit from the parsimonious specification is the possibility to estimate consumption functions also for each country separately. For the employed panel stationarity tests for example this is a prerequisite, as they are based on country specific ADF-tests.

\(^2\) Like the role of interest rates (which did not appear relevant in preliminary specification testing), the degree of financial liberalization or structural breaks.
to investigate real estate wealth effects. The framed expression contains the rental housing market variables. An implicit assumption underlying formulation (1) is that the rental rate elasticity varies linearly with the proportion of renters. This is reflected in the use of the composed variable $S_{it}R_{it}$ (rather than $R_{it}$ alone), which makes the rental rate elasticity of consumption ($=\alpha_5 S_{it}$) countryspecific. The term $\alpha_5$ itself can therefore be interpreted as the rental rate elasticity for a fictitious country with a renter share of 1.

Rental rates $R_{it}$ are only available as indices. Therefore, they need to be rescaled by unknown countryspecific factors to get rental rates comparable across countries.$^3$ These factors are part of the unobserved country-specific component $\nu_i$.\(^4\) Furthermore these factors also enter the coefficient $\alpha_6i$ of the renter share $S_{it}$, which, therefore, must be modeled country specific in (1). It should be noted, that formulation 1 entails the possibility for opposite effects of renter share and rental rate.

Apart from country specific effects formulation (1) also contains time specific effects $\theta_t$, which are to be estimated jointly with the other coefficients. These time specific effects serve to remove a simple kind of potential cross country correlation from the residuals $\varepsilon_{it}$.

Equation (1) will be referred to as the shortrun equation, but it will be used here also to derive longrun elasticities implicitly given by the estimated short run coefficients. Direct estimation of longrun elasticities will be based instead on the cointegrating relationship

$$\tilde{C}_{it} = \beta_1 \tilde{Y}_{it} + \beta_2 \tilde{P}^H_{it} + \beta_3 \tilde{P}^S_{it} + \beta_4 \tilde{S}_{it} \tilde{R}_{it} + \beta_5 i \tilde{S}_{it} + \nu_i + \varepsilon_{it},$$

based on demeaned data $\tilde{X}_{it} = X_{it} - \frac{1}{n} \sum_{j=1}^{n} X_{jt}$ for $X = \{C, Y, SR, S, P^H, P^S\}$. Equation (2) will be referred to as the longrun equation. Estimation of the coefficients in (2) will in fact be based on more extensive regression equations including other variables as well (see below).

$^3$ The underlying non-log form (indicated with bars) of the relationship, including such countryspecific rescaling factors $b_i$ for comparability of rent indices across countries, would be $\tilde{C}_{it} = \ldots (b_i R_{it})^\alpha S_{it} \ldots$ Taking logs thereof makes the renter share a countryspecific variable as in 1.

$^4$ Analogous reasoning also applies to the price variables $P^H_{it}$ and $P^S_{it}$, whose corresponding rescaling coefficients are also lumped together in the country-specific effect $\nu_i$. 

---

3 The underlying non-log form (indicated with bars) of the relationship, including such countryspecific rescaling factors $b_i$ for comparability of rent indices across countries, would be $\tilde{C}_{it} = \ldots (b_i R_{it})^\alpha S_{it} \ldots$ Taking logs thereof makes the renter share a countryspecific variable as in 1.

4 Analogous reasoning also applies to the price variables $P^H_{it}$ and $P^S_{it}$, whose corresponding rescaling coefficients are also lumped together in the country-specific effect $\nu_i$. 

---

3
2.2 Econometric Issues

Given the low numbers of observations in longitudinal and cross-section dimension of the present data set, the use of panel techniques based on pooling of data promises more reliable estimates than pure cross section or time series techniques. The availability of panel data furthermore alleviates problems arising from non-stationarity of data in a pure time series context (see e.g. Kao [1999] or Phillips & Moon [2000]). The latter fact motivates comparison of the magnitudes of implicit long run elasticities from the short run equation (1) with direct long run elasticity estimates from (2).

To check the sensitivity of the results regarding the choice of estimators, different techniques are employed to estimate each of (1) and (2). For (1) these are a bias corrected version of the Least Squares Dummy Variables approach (LSDVC) proposed in Kiviet [1995] and the system version of the generalized method of moments estimator (SYSGMM) of Arellano & Bover [1995] and Blundell & Bond [1998]. Encouraging comparisons of the finite sample properties of these estimators with possible others can be found in Judson & Owen [1999] or Lusinyan [2005].

To estimate (2) Dynamic OLS for panel data (DOLS, see Kao & Chiang [2000] and Mark & Sul [2002]) and the Pooled Mean Group (PMG) estimator by Pesaran et al. [1995] are used. DOLS has been applied in the relevant literature for example in Dvornak & Kohler [2003] or Catte et al. [2004], the PMG approach for example in Barrell & Davis [2004] and Ludwig & Sløk [2004]. For both these approaches equation (2) only serves to define a corresponding error correction term, while the equations actually estimated include the differenced variables and eventually lags and leads thereof. Furthermore both these approaches estimate country-specific short run dynamics, which can capture different adjustment speeds towards long run equilibrium for each country.

To test for unit roots two panel based approaches are used. One is the Fisher-test suggested in Maddala & Wu [1999]. The other was proposed in Im et al. [2003] (IPS) as t-bar statistic. Both of these commonly used tests are based on country specific unit-root tests, but require prior removal of common effects (demeaning). Consequently regression
(2) was also formulated for demeaned data rather than with explicit time effects. Comparisons of the finite sample properties of unit-root testing approaches in Banerjee [1999], Baltagi et al. [2000] or Choi [2001], indicate that the chosen tests have more power than potential alternatives for the dimensionality of the panel available.

Cointegration of the variables in equation (2) is tested with the non-parametric group $t$-test proposed in Pedroni [1999] and the Mean Variance Ratio test of Westerlund [2005]. These panel approaches share the assumption that all variables are I(1) and the cointegrating relationship is unique. Rank based tests of the Johansen variety to check this uniqueness could not be applied due to limited availability of longitudinal data. A further assumption is that no subgroups of variables are themselves cointegrated. This may be suspected particularly to hold for house prices and rents. But the evidence for such cointegration is mixed (see Gallin [2004]). For the composed variable $S_{it}R_{it}$ used in the present paper (rather than $R_{it}$) the cointegration tests rejected this hypothesis. This was also the case for income and house prices, in line with Gallin [2006], who investigated US data for several metropolitan areas.

3 Data

The results to be presented are based on aggregate yearly data for 18 OECD countries observed between 1970 and 2004. For the bulk of countries complete data were available only for the 1980-2004 period, making the panel rather unbalanced.

Rental price indices for the present analysis are drawn primarily from the data base of the International Labour Organization (ILO), which reports the figures from regular national collections of consumer prices conducted through the national statistical offices. These indices in most cases reflect average rental costs per square meter paid by all tenants. For Australia, Austria, Ireland and Norway instead the ILO figures correspond to the broader category housing, which includes such items as fuel and light.

---

5 See for example Pedroni [1999] on explicit estimation of time effects vs. demeaning.
6 Australia, Austria, Belgium, Canada, Denmark, Finland, France, Germany, Ireland, Italy, Japan, Netherlands, New Zealand, Norway, Spain, Sweden, UK and USA. For lack of relevant data other countries like Switzerland or Portugal had to be excluded.
7 The exceptions are Germany, France, UK and the USA, for which longer series are available.
8 For a detailed, country by country description of the ILO rent indices see http://laborsta.ilo.org
Table 1: Annual % change of real rents (averages)

<table>
<thead>
<tr>
<th>Country</th>
<th>70-80</th>
<th>81-90</th>
<th>91-00</th>
<th>00-04</th>
<th>70-80</th>
<th>81-90</th>
<th>91-00</th>
<th>00-04</th>
</tr>
</thead>
<tbody>
<tr>
<td>Australia</td>
<td>0.23</td>
<td>1.31</td>
<td>-1.79</td>
<td>0.94</td>
<td>Italy</td>
<td>-2.92</td>
<td>1.47</td>
<td>2.66</td>
</tr>
<tr>
<td>Austria</td>
<td>2.97</td>
<td>2.98</td>
<td>2.42</td>
<td>1.98</td>
<td>Japan</td>
<td>-0.89</td>
<td>1.35</td>
<td>0.88</td>
</tr>
<tr>
<td>Belgium</td>
<td>1.64</td>
<td>0.83</td>
<td>1.14</td>
<td>0.24</td>
<td>NL</td>
<td>0.62</td>
<td>2.56</td>
<td>2.31</td>
</tr>
<tr>
<td>Canada</td>
<td>-3.75</td>
<td>-0.03</td>
<td>-0.06</td>
<td>-0.80</td>
<td>NZ</td>
<td>-1.52</td>
<td>2.66</td>
<td>2.27</td>
</tr>
<tr>
<td>Denmark</td>
<td>-0.62</td>
<td>1.27</td>
<td>0.85</td>
<td>0.78</td>
<td>Norway</td>
<td>-1.25</td>
<td>0.39</td>
<td>0.35</td>
</tr>
<tr>
<td>Finland</td>
<td>-1.06</td>
<td>-0.61</td>
<td>1.45</td>
<td>1.24</td>
<td>Spain</td>
<td>-3.69</td>
<td>-2.00</td>
<td>1.96</td>
</tr>
<tr>
<td>France</td>
<td>-1.15</td>
<td>1.23</td>
<td>0.80</td>
<td>0.08</td>
<td>Sweden</td>
<td>-2.38</td>
<td>1.71</td>
<td>1.61</td>
</tr>
<tr>
<td>Germany</td>
<td>1.13</td>
<td>1.75</td>
<td>2.10</td>
<td>0.33</td>
<td>UK</td>
<td>0.62</td>
<td>3.26</td>
<td>2.79</td>
</tr>
<tr>
<td>Ireland</td>
<td>-5.59</td>
<td>-0.35</td>
<td>2.01</td>
<td>0.13</td>
<td>USA</td>
<td>-0.04</td>
<td>1.37</td>
<td>0.25</td>
</tr>
</tbody>
</table>

Nominal rent indices are deflated with the Harmonized CPI except for AU, CA, JP, NZ and USA, where the standard CPI was used instead. Sources for nominal rent indices: ILO, national statistical offices and others as mentioned in the text; Source for (H)CPI: OECD Economic Outlook.

These and other shortcomings of the ILO data are addressed by adaptations based on additional sources. For Austria, Ireland and Norway price indices for rents proper, provided by the respective National Statistical Office, were used instead of ILO data. No such correction was possible for Australia. The figures for the USA are drawn from Crone et al. [2004], who calculated rental series corrected for non response and ageing bias. The data for Germany are taken from Frick [2006] and linked with a corresponding long series from the National Statistical Office Wiesbaden for 1970-1985. Statistics New Zealand and Statistics Sweden provided consistently linked series whereas the corresponding ILO series contain breaks due to changes in price collection practice. Yearly average growth rates of the resulting rent indices after deflation are provided in Table 1 for each country. As can be seen, these growth rates differ widely between time periods and countries.

Renter shares are collected from various sources. The Royal Institute of Chartered Surveyors compiles long series for the UK and France, the German Statistical Office provides such data for Germany and the US Census Bureau for the USA. For other countries I relied on data collected by the Danish Agency for Enterprising and Housing and figures reported in ECB [2003, p. 26], Catte et al. [2004b] and Arévalo & Ruiz-Castillo [2006]. In some cases series from different sources had to be linked to extend the observation period to at least 1980-2004. In general renter shares for earlier time periods are available only at larger intervals. In these cases missing values were linearly interpolated as for example
in Ludwig & Sløk [2004]. With these interpolations the first differences of renter shares contain little information and therefore are skipped as eventual auxiliary regressors in the estimation procedures. Besides “Owners” and “Renters” relevant sources occasionally list additional types of tenure. This might explain differences in stated “renter shares” between the sources mentioned above for individual countries. In conflicting cases the ultimate criterion for choosing between sources was maximum length of observation period. An overview of the series constructed in this manner is given in Table 2.

The data set about rental housing compiled in this fashion is still far from ideal regarding cross-country comparability and consistency over longer time horizons. This reflects the often poor coverage of rental housing in national statistics, particularly for earlier periods and may explain, why so little empirical research on this topic can be found in the literature.

The other variables used in this analysis are standard in studies on real estate wealth effects. These variables include household consumption, household disposable income, consumer prices, stock prices and the instrumental variables governmental expenditures, interest rates and population, which were all drawn from the OECD data base. For estimation purposes consumption and income are expressed on a per capita basis. Time series on real house prices are from the Bank for International Settlements (BIS) as in most other relevant studies.\(^9\)

---

\(^9\) These prices are the inflation adjusted residential property prices from BIS calculations based on
Income is measured as real household disposable income as reported by the OECD. Wealth effect studies occasionally use wage and salary income instead to avoid potential problems with double counting of property income. But control estimates with this alternative income measure did not yield much different results. For better comparability with the bulk of relevant literature therefore real household disposable income is used instead. The nominal variables are deflated by the harmonized consumer price index (HCPI) where available (for all European countries), else with the standard CPI. The HCPI is preferable, because it does not include imputed rents, which cause problems regarding CPI comparability across countries.

4 Results

4.1 Shortrun estimation

Apart from the appearance of lagged consumption as regressor, the main econometric problem with equation 1 is potential endogeneity, particularly of the income variable. To cope with this problem instrumental variables for income, rental rate, house prices and stock prices are used in both estimation approaches. The renter share instead was tested positively for exogeneity. As instruments the lagged values of the seemingly endogenous variables were used plus government expenditures, short term interest rates and population growth.

SYSGMM addresses endogeneity by construction and in a more general fashion by making use of all lagged values of all variables as instruments. But in line with other studies using SYSGMM the lag number was restricted here to avoid overidentification. This could be achieved reliably by using only the first three lags. The test for AR(2) errors could not be rejected (as it should) at standard confidence levels. Therefore an equation with an additional lag of the consumption variable was estimated for comparison. While this removed AR(2) errors, the coefficient estimates remained much the same. For better comparability with the bulk of the real estate wealth effect literature therefore only results national data.
for the single lag specification are given here. The standard errors reported in table 3 for the longrun elasticities (derived from shortrun coefficients) are the asymptotic ones.

<table>
<thead>
<tr>
<th></th>
<th>(C_{i,t-1})</th>
<th>(Y_{i,t})</th>
<th>(S_{i,t}R_{i,t})</th>
<th>(P_{i,t}^H)</th>
<th>(P_{i,t}^S)</th>
</tr>
</thead>
<tbody>
<tr>
<td><strong>LSDVC</strong></td>
<td>short</td>
<td>0.773</td>
<td>0.169</td>
<td>-0.117</td>
<td>0.015</td>
</tr>
<tr>
<td></td>
<td>SE</td>
<td>0.034</td>
<td>0.022</td>
<td>0.035</td>
<td>0.006</td>
</tr>
<tr>
<td></td>
<td>long</td>
<td>0.747</td>
<td>-0.517</td>
<td>0.065</td>
<td>0.095</td>
</tr>
<tr>
<td></td>
<td>SE</td>
<td>0.119</td>
<td>0.139</td>
<td>0.024</td>
<td>0.020</td>
</tr>
<tr>
<td><strong>SYSGMM</strong></td>
<td>short</td>
<td>0.685</td>
<td>0.236</td>
<td>-0.161</td>
<td>0.028</td>
</tr>
<tr>
<td></td>
<td>SE</td>
<td>0.057</td>
<td>0.041</td>
<td>0.049</td>
<td>0.013</td>
</tr>
<tr>
<td></td>
<td>long</td>
<td>0.751</td>
<td>-0.512</td>
<td>0.091</td>
<td>0.046</td>
</tr>
<tr>
<td></td>
<td>SE</td>
<td>0.119</td>
<td>0.152</td>
<td>0.030</td>
<td>0.014</td>
</tr>
</tbody>
</table>

Table 3: Shortrun Estimates

The main result from Table 3 given in column \(S_{i,t}R_{i,t}\) is that the rental rate elasticities of aggregate consumption are significantly negative. The shortrun estimates for a fictitious country with renter share of 1 are -0.117 and -0.161 for LSDVC and SYSGMM respectively. They indicate, that for a renter share of 32% as in the EU-25 or the USA, the short run rental rate elasticity of aggregate the rental rate roughly speaking is -0.045. The derived longrun estimates instead are very similar (-0.517 resp. -0.512). This implies, for Germany for example with a renter share of about 51% \((S_{it} = 0.51)\), a long run elasticity of about -0.26. The actual increase of real rental rates in Germany between 1994-2004 amounted to 11.9%, implying a loss of yearly consumption of about 3% compared with the fictitious constant real rent case since 1994. The analogous loss for the USA with roughly 31% renters and an actual 8.4% real rent increase in the same period amounts to about 1.4%.

Regarding the elasticities of the two wealth proxies (stock prices and house prices) the long run figures are in the 0.05 – 0.10 region and significant at usual confidence levels. Estimated with LSDVC, the stock wealth effect appears stronger than the housing wealth effect, with SYSGMM it is the other way round. Both of these findings, significant effects in the cited range on one hand and unclear relative importance of the two wealth components on the other, are broadly in line with the findings in the relevant literature on real estate wealth effects. Thus, adding a rental housing segment to the consumption equation does not seem to alter the basic conclusion from this strand of literature. But, as discussed in section 4.3, the joint consideration of rental rate and house price elasticities
yields a somewhat different assessment.

4.2 Longrun estimation

Stationarity, Cointegration

The country specific augmented Dickey-Fuller tests for stationarity yielded mixed results for all variables. But due to the shortness of some of the series these tests are likely to have little power (see eg. Enders [1995]) and therefore give little guidance.\textsuperscript{10} Stationarity tests based on panel data promise more power and, indeed, yielded less ambiguous results. The first of these tests, the Fisher-Test, was applied to the unbalanced panel of all countries. The second, the IPS test, could be carried out only for two subsamples: “France, Germany, UK, USA 1970-2004” and “All countries 1980-2004”, because it requires balanced panels. Based on these panel tests the Null-hypothesis of non-stationarity can not be rejected for consumption and income at usual significance levels. Also for the composed variable $S_{i,t} R_{i,t}$ and for $S_{i,t}$ itself these tests point towards I(1) rather than stationarity. For the two wealth proxies (stock prices and house prices) instead all tests reject the unit root hypothesis. Particularly for stock prices this seems to contradict usual wisdom, but it should be recalled that the tests apply to the demeaned series here. Furthermore, given the construction of the tests, these findings do not imply that $P^H$ and $P^S$ are stationary for all countries, but only for a significant proportion thereof. Nevertheless this stationarity evidence, strictly speaking, calls for exclusion of these two price variables in the cointegration tests to follow, because they require I(1) variables.

Given this objections the cointegration tests were carried out separately for three groups of variables: Baseline specification A includes all variables used also in estimating the short run dynamics ($C, Y, SR, S, P^S, P^H$). Specification B includes only ($C, Y, SR, S$) because for house and stock prices no clear I(1) evidence was found. In specification C ($C, Y, P^S, P^H$) finally the two rental housing market related variables are excluded. This serves to check whether cointegration evidence for a typical variable set encountered in the

\textsuperscript{10} The maximum lag length considered for these tests was 9 and the choice of the actual lag length was based on the Schwartz-Bayes information criterion. Of the original 24-34 observations therefore only 15-25 could be used for testing.
real estate wealth effect literature is any better than the one for the baseline specification A of this paper.

Both cointegration tests used, Pedroni’s $t$-test and Westerlund’s mean variance ratio test, require symmetric panels. Therefore only the subsamples “All countries 1980-2004” (=S1) and “France, Germany, UK, USA 1970-2004” (=S2) were tested. With S1 cointegration evidence is found for specifications A (with $p$-values of 0.01 and 0.03 for Pedroni’s and Westerlund’s tests respectively) and B (0.10 and 0.02), but not for specification C ($p=0.46$ and 0.70). For S2 instead no test indicates cointegration for any of A, B or C ($p$-values $\geq 0.20$), but again specification C does worst ($p$-values $\geq 0.86$). Overall the cointegration evidence therefore is mixed, but in any case is better for specifications including rental market variables than for the specification without them.\textsuperscript{11}

**Estimation results**

Results from the estimation of longrun (cointegration) relationships are reported in Table 4 for the baseline specification A of variables.\textsuperscript{12} These results lend themselves for comparison with the results from the shortrun equation reported in Table 3. This implies the inclusion of variables $P^H$ and $P^S$, which in their demeaned form, as mentioned above, appear as I(0). Thus, strictly speaking the cointegration evidence for this specification A from above is not applicable. Estimation with DOLS is restricted to symmetrical panels and therefore is carried out separately for the two subsamples S1 and S2 already used in cointegration testing. Due to limited time series information only 2 lags and 1 lead of the differenced variables are used for the endogeneity correction via DOLS, like in Mark & Sul [2002]. The long run estimates from the Pooled Mean Group (PMG) approach are based on the full, unbalanced set of observations.

\textsuperscript{11} Additional testing for cointegration was carried out based on the bounds approach of Pesaran et al. [2001]. Given the evidence from the unit-root tests, this approach would be more appropriate particularly for specifications A and C, because it is valid also for mixtures of I(0) and I(1) variables. Due to data limitations this type of testing could only be carried out for a maximum of 3 lags and specifications B and C, but the general picture emerging supports the findings above: Cointegration evidence for specification B is better than for specification C.

\textsuperscript{12} Additional estimation for specification B (without the wealth variables), for which cointegration evidence was best, lead to significantly negative rental rates elasticities in the -0.35 range (depending on the subsample used).
The figures reported in Table 4 confirm the basic findings from section 4.1. The rental rate elasticity of consumption is again negative according to both approaches used and highly significant in two of the three cases considered. For a fictitious country with $S_{i,t}=1$ this elasticity is between $-0.42$ and $-0.53$, depending on the subsample and similar to the corresponding figures derived from estimates of the shortrun equations. In particular the $-0.48$ elasticity from the PMG approach, based on the same sample as the shortrun estimates, comes close to the $-0.51$ figures from the latter. Table 4 furthermore shows that stock wealth effects remain significantly positive, while housing wealth is significantly positive in two cases and insignificantly negative in a third. Thus, at least the results for the rental rate elasticity are fairly robust with regard to the choice of the estimation technique. This strengthens the case for inclusion of rental housing variables into the consumption equation compared to the usual inclusion of wealth variables.

### 4.3 House prices and rental rates

The user cost approach to housing (e.g. Miles [1994]) suggests, that rental rates and house prices are linked through a portfolio equilibrium condition like $R = (i + d)P^H$. In this formulation $i$ and $d$ denote interest and depreciation rate respectively and taxes and expected capital gains are ignored for simplicity. This would imply that for fixed interest and depreciation rates house prices move proportionally with rental rates. In this case the correct rental rate elasticity for country $i$ would have to be calculated as sum $\alpha_3 + \alpha_5 S_{it}$ (rather than $\alpha_5 S_{it}$). These elasticities with associated standard errors are reported in Table 5 for different renter shares $S_{it}$.
The short run elasticity estimates for a country with 32% renter share (like the USA or the typical EU-25 country) would then for example be -0.022 and -0.024 for LSDVC and SYSGMM respectively. Analogously the derived longrun elasticities would be -0.100 and -0.073 respectively. But none of these figures appears significantly different from zero at usual confidence levels, albeit achieved significance levels increase with the renter share. It must be noted, though, that by the same logic the above results would also imply that the real estate wealth effect upon consumption is insignificant.13 This contradicts the evidence from a large body of literature where these effects were estimated to be significantly positive. Together with the fact, that rental housing variables appear statistically relevant for explaining consumption, this renders the relevance of this condition in the present context questionable. The appearance of time varying interest rates in this portfolio equilibrium condition (and time varying expected capital gains in a more elaborate version thereof) casts additional doubts on the simple proportionality between house prices and rental rates. In particular if there are any causal relationships between rental rates (and/or house prices) and interest rates (and/or capital gains) the above used elasticity measure could be highly misleading. The low (cross country average) correlation between these variables of 0.12 (with a standard deviation of 0.46) substantiate these objections. Thus, invoking such proportionality to calculate rental rate elasticities seems rather questionable.

13 In Al-Eyd et al. [2005] such a potentially offsetting impact from rental rates for housing wealth effects was hinted at.

---

Table 5: Rental rate elasticities of consumption when \((\partial P^H / \partial R) \times (R/P^H) = 1\)

<table>
<thead>
<tr>
<th></th>
<th>shortrun</th>
<th></th>
<th></th>
<th>longrun</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>S_{it} =</td>
<td>0.20</td>
<td>0.32</td>
<td>0.40</td>
</tr>
<tr>
<td>LSDVC</td>
<td>-0.008</td>
<td>-0.022</td>
<td>-0.032</td>
<td>-0.038</td>
</tr>
<tr>
<td>SE</td>
<td>0.010</td>
<td>0.014</td>
<td>0.017</td>
<td>0.047</td>
</tr>
<tr>
<td>SYSGMM</td>
<td>-0.004</td>
<td>-0.024</td>
<td>-0.036</td>
<td>-0.011</td>
</tr>
<tr>
<td>SE</td>
<td>0.013</td>
<td>0.017</td>
<td>0.020</td>
<td>0.049</td>
</tr>
</tbody>
</table>

The short run elasticity estimates for a country with 32% renter share (like the USA or the typical EU-25 country) would then for example be -0.022 and -0.024 for LSDVC and SYSGMM respectively. Analogously the derived longrun elasticities would be -0.100 and -0.073 respectively. But none of these figures appears significantly different from zero at usual confidence levels, albeit achieved significance levels increase with the renter share. It must be noted, though, that by the same logic the above results would also imply that the real estate wealth effect upon consumption is insignificant. This contradicts the evidence from a large body of literature where these effects were estimated to be significantly positive. Together with the fact, that rental housing variables appear statistically relevant for explaining consumption, this renders the relevance of this condition in the present context questionable. The appearance of time varying interest rates in this portfolio equilibrium condition (and time varying expected capital gains in a more elaborate version thereof) casts additional doubts on the simple proportionality between house prices and rental rates. In particular if there are any causal relationships between rental rates (and/or house prices) and interest rates (and/or capital gains) the above used elasticity measure could be highly misleading. The low (cross country average) correlation between these variables of 0.12 (with a standard deviation of 0.46) substantiate these objections. Thus, invoking such proportionality to calculate rental rate elasticities seems rather questionable.

13 In Al-Eyd et al. [2005] such a potentially offsetting impact from rental rates for housing wealth effects was hinted at.
5 Conclusions

This paper estimates the impact of changing housing rental rates upon aggregate consumption. The starting point is the standard formulation of a dynamic consumption equation as used in the real estate wealth effect literature. This formulation is extended to include the housing rental rate and the renter share as additional explanatory variables. Because of very different renter shares across countries the elasticity estimates are modeled as country specific. To estimate the relevant consumption equation yearly data for 18 OECD countries covering the period from 1970 to 2004 are used. Several panel regression techniques are employed to come up with long run elasticity estimates derived indirectly from short run dynamic equations and directly from long run cointegrating relationships. The results appear robust across estimation techniques and indicate a significantly negative impact of increasing rental rates upon aggregate consumption. The relevant long run elasticities for a fictitious country with 100% renters are in the range of -0.42 to -0.53 for the various techniques used. This implies for example for the EU-25 or the USA, with renter shares of around 32%, that a 1% increase of the real rental rate lowers aggregate consumption between -0.13% and -0.17% in the long run. For the actual 11.9% increase of real rents in Germany between 1994-2004 and a 51% rentershare, one can calculate a loss of yearly consumption of around 2.8% compared with the fictitious case of constant real rents since 1994. The comparable loss for the USA with an actual 8.4% real rent increase in the same period and a 31% rentershare amounts to about 1.2%.

The estimated elasticities of the two wealth components (financial wealth and housing wealth) remain significantly positive despite the inclusion of the rental housing market variables and are roughly in the 0.05 - 0.10 range known from previous empirical studies. These results substantiate the case for monitoring rental rates more closely when aggregate consumption demand is at stake. Unfortunately satisfactory data describing conditions of rental housing markets over longer time horizons or across countries are still hard to find. But the recent attempts of national statistical offices to provide better figures will help to improve the preliminary evidence on this topic found in the present study.
References


Boone, L. & Girouard, N. [2002], The stock market, the housing market and consumer behaviour, OECD Economic Studies 35, 2002/2, pp. 175-200, OECD.


Catte, P., Girouard, N., Price, R. & André, C. [2004a], The contribution of housing markets to cyclical resilience, OECD Economic Studies 38, OECD.

Catte, P., Girouard, N., Price, R. & André, C. [2004b], Housing markets, wealth and the business cycle, Economic department working papers no. 394, OECD.

Choi, I. [2001], ‘Unit root test for panel data’, *Journal of International Money and Banking* 20, 249–272.


Lusinyan, L. [2005], Dynamic panel data models in macroeconomic applications, Presentation at the 12th conference on panel data, basle/ch.


