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Trend and cycle features in German residential investment before and after reunification

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Abstract

Real residential investment in Germany is found to be cointegrated with population, real national income per capita and real house prices. This evidence is consistent with a model where the trend in housing demand is determined by demographic factors and economic well-being to which supply adjusts so slowly that real house prices are affected persistently. Reunification seems to have induced two structural changes in the empirical housing market model. First, the speed of equilibrium adjustment via residential investment slowed down substantially and real house prices lost the capacity to contribute to the adjustment process. Second, the degree of persistence in the error correction term increased a lot. The changing features are key to explain significant differences in alternative trend-cycle decompositions of residential investment.

Keywords: Residential investment, vector autoregression, trend-cycle decomposition, Germany.

JEL classification: E22, C32.

Non-technical summary

In this paper, real residential investment in Germany over the last 35 years is modelled with a particular focus on the role of trend and cyclical factors. This task is complicated by the peculiarities of the German reunification because they have influenced housing market developments to a significant degree. In the early 1990s, the existing supply of dwellings in western Germany did not meet the demand for housing, which rose sharply especially as a result of immigration. Moreover, the residential stock in eastern Germany was run-down and had to be renovated. Dwellings construction expanded massively between 1991 and 1995. Having remained elevated until the year 2000, residential investment declined for a lenghty period which did not end until 2005 when it turned into a moderate cyclical recovery. The housing market was thus characterized by a boom-bust cycle in residential investment, while house prices changed only marginally in nominal terms and decreased fairly steadily in real terms.

Real residential investment is modelled within a cointegrated vector autoregression together with population, real income per capita and real house prices. The cointegrating relation which is established between the four variables can be interpreted as a long-run equilibrium relationship on the housing market. The trend in housing demand is described by demographics and per-capita income. The adjustment of dwellings supply via residential investment is rather slow, implying that lasting tensions on the housing market are reflected in persistent house price movements.

As the sample comprises data before and after the reunification, interesting comparisons can be made and changes in the model structure can be uncovered. The pronounced up-and-down movement of residential investment after reunification is reflected in the long-run residuals of the cointegrating relation, as its time series is characterized by an increase in the degree of persistence since 1991. Furthermore, one can conclude that since 1991 house prices have contributed less than before to the clearing of demand and supply on the dwellings market.

In the given model setup, the trend-cycle decomposition of the Gonzalo and Granger (1995) type yields the following results. Composed as a linear combination of the three weakly exogenous variables, the trend component of residential investment is rather smooth. As the long-run residual series constitutes the cycle component, the pronounced boom-bust movement is regarded as a cyclical phenomenon. In this respect, the Gonzalo-Granger decomposition differs substantively from the results of univariate statistical approaches to trend extraction such as the Hodrick-Prescott and the Baxter-King filters.

Nicht technische Zusammenfassung

In diesem Papier werden die realen Wohnungsbauinvestitionen in Deutschland über die letzten 35 Jahre modelliert, wobei insbesondere die Rolle von Trend und zyklischen Faktoren berücksichtigt wird. Diese Aufgabe wird durch die Besonderheiten im Umfeld der deutschen Wiedervereinigung erschwert, weil sie die Entwicklung des Häusermarkts über einen beträchtlichen Teil des Untersuchungszeitraums geprägt haben. Anfang der neunziger Jahre stand der insbesondere durch Migration stark gestiegenen Nachfrage nach Wohnraum in Westdeutschland kein entsprechendes Angebot gegenüber. Außerdem musste der desolate Immobilienbestand in Ostdeutschland erneuert werden. Zwischen 1991 und 1995 wurde der Wohnungsbau daher stark ausgeweitet. Nachdem er etwa bis zum Jahr 2000 auf erhöhtem Niveau verblieben war, kam es zu einer lang anhaltenden Korrekturphase, die erst nach 2005 in eine moderate zyklische Erholung überging. Der Wohnungsmarkt erlebte damit einen sehr ausgeprägten Investitionszyklus, während sich die Häuserpreise in dieser Zeit nominal kaum veränderten und in realer Betrachtung sogar recht stetig gefallen sind.

Die realen Wohnungsbauinvestitionen werden zusammen mit der Wohnbevölkerung, dem realen Pro-Kopf-Einkommen und den realen Häuserpreisen mittels einer kointegrierten Vektorautoregression modelliert. Die Kointegrationsbeziehung, die sich zwischen diesen vier Variablen etablieren lässt, kann als eine langfristige Gleichgewichtsbeziehung auf dem Häusermarkt interpretiert werden. Die Nachfrage nach Wohnraum wird im Trend durch die demographische Entwicklung und das Pro-Kopf-Einkommen beschrieben. Die Anpassungen des Angebots über Investitionen erfolgen recht langsam, sodass sich anhaltende Spannungen auf dem Wohnungsmarkt in persistenten Preisbewegungen niederschlagen.

Da der Stützzeitraum Daten aus der Zeit vor und nach der Wiedervereinigung einschließt, lassen sich interessante Vergleiche anstellen und Veränderungen in der Modellstruktur aufdecken. Die ausgeprägte Auf- und Abwärtsbewegung der Wohnungsbauinvestitionen nach der Wiedervereinigung spiegelt sich in den Langfristresiduen der Kointegrationsbeziehung wider, deren Zeitreihe ab 1991 eine deutlich erhöhte Persistenz aufweist. Ferner kann geschlossen werden, dass die Häuserpreise nach 1990 weniger als zuvor dafür gesorgt haben, dass Angebot und Nachfrage ins Gleichgewicht kommen.

Die Trend-Zyklus-Zerlegung nach Gonzalo und Granger (1995) kommt in der vorliegenden Modellstruktur für die Wohnungsbauinvestitionen zu den folgenden Resultaten: Als Linearkombination der drei schwach exogenen Variablen verläuft die Trendkomponente vor und nach der Wiedervereinigung relativ glatt. Da die Langfristresiduen der Kointegrationsbeziehung die Zykluskomponente bilden, wird der ausgeprägte Auf- und Abschwung der Wohnungsbauinvestitionen zwischen 1991 und 2005 als zyklisches Phänomen aufgefasst. Hierin unterscheidet sich die Gonzalo/Granger-Zerlegung deutlich von den Ergebnissen univariater statistischer Trendextraktionsverfahren wie dem Hodrick/Prescott- und dem Baxter/King-Filter.

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1 Introduction

The evolution of the German housing market has differed from that in many other industrialized countries for quite some time. While house prices recently underwent a pronounced up-and-down movement in the U.S. and some western European countries, they remained flat in Germany. Furthermore, weak residential investment has steadily weighed on economic growth in Germany since the mid-1990s, whereas the construction of new dwellings had been a major stimulus for economic growth elsewhere until the house price bubbles, which had often emerged simultaneously, ultimately burst.

A boom-bust movement in residential construction happened in Germany about ten years earlier.[2] After the fall of the iron curtain, strong immigration into western Germany and the need to improve the overall run-down housing stock in eastern Germany triggered a steep upswing in housing investment which was additionally bolstered by exuberant government promotion. Towards the end of the 1990s, market conditions not only relaxed but also turned, at least in part, into a situation of misallocation and over-investment (Deutsche Bundesbank, 2002, for instance). As a consequence, the government gradually curtailed the quite advantageous depreciation allowances for investors as well as the subsidies and grants offered to homebuyers. Against the backdrop of a stagnant population trend and gloomy income prospects, demand for new dwellings then declined for several years. Until recently, residential investment has been rather weak, with some stimuli created by modernization activity in the existing housing stock.

In applied macroeconomic research in Germany, residential investment has attracted little interest in recent years. One reason for the lack of attention might be the fact that, in this series, the post-reunification period did not end until the middle of the past decade, making it difficult to separate the time series properties of "normal" phases from the special pattern induced by the circumstances in the early 1990s. Looking at the complete boom-bust movement, this paper identifies what has remained unchanged since the west German era and what has changed since then. On the one hand, the changes are temporary insofar as they are attributed to this seminal event, which caused a big shock and triggered

[1]Corresponding address: Deutsche Bundesbank, Economics Department, Wilhelm-Epstein-Str. 14, 60431 Frankfurt am Main, Germany, email: thomas.knetsch@bundesbank.de. The paper has been written within the joint research project (JRP) "The macroeconomics of housing markets" organized by the Deutsche Bundesbank, Banque de France, Banca d'Italia and Banco de España. Valuable comments by Luis Álvarez, Andrew Benito, Laurent Ferrara and Karl-Heinz Tödter are gratefully acknowledged. The paper expresses the author's personal opinion which does not necessarily reflect the views of the Deutsche Bundesbank.

[2]It is worth mentioning that the boom-bust cycle in Germany was first and foremost a quantity phenomenon. In the boom phase between 1990 and 1995,house prices in western Germany rose less strongly compared with the corresponding episodes in other countries. Thereafter, they remained more or less flat instead of falling significantly.

adjustment processes thereafter. On the other hand, reunification might also have brought the housing market to a new long-run equilibrium.

The boom-bust movement which appeared in the housing market after reunification changed the time series properties of residential investment which had been manifest in the west German era. In general, simple statistical trend-cycle filter techniques are not able to identify the post-reunification movement as a cycle of different duration and amplitude but tend to assign it to the trend component. This paper, however, suggests a model which comes out with a rather smooth trend in residential investment driven by economic fundamentals such as demographics, economic well-being and house price developments. In addition, it figures out that the forces of equilibrium correction have weakened substantively in the German housing market since reunification. The speed of adjustment in residential investment has slowed down significantly and house prices seem to have lost the capacity to equilibrate demand and supply in this market. Moreover, the evidence suggests a marked increase in the degree of persistence in the cyclical component of residential investment.

It is as yet an open question whether, or to what extent, the changed features will return to the initial patterns in the future. The paper aims at presenting facts on aggregate housing market developments in Germany over the last 35 years, including a big shock approximately in the middle of this time period. However, it will not address the implications these observations may have on forecasting and structural analysis.

In the first part of the paper, a cointegration analysis yields an overall stable long-run economic relationship between residential investment, population, real national income per capita and real house prices. It is worth mentioning that this structure deviates from the consumption good hypothesis, which is the theoretical concept guiding the modelling of residential investment in many macroeconometric models.[3] The second part is devoted to the major structural change of the early 1990s, namely a substantial deceleration in the speed of adjustment to the long-run equilibrium relationship, suggesting that residential investment displays a Gonzalo and Granger (1995) cycle component of higher persistence and amplitude.[4] The pitfalls this structural change implies for the application of standard statistical trend-cycle filters are highlighted in addition. In the final section, some conclusions are drawn and the limitations of the present analysis are mentioned.

2 Long-run determinants of residential investment

In the first part, the theoretical foundations of residential investment are discussed. As a microeconomic approach based on utility theory suffers from the fact that representative behavior is virtually indetectable in the aggregate housing market owing to various forms

[3]Carnot et al. (2005) give a brief overview of the general modelling principles of residential investment in macroeconometric models. See Heilemann (2004) for an implementation in a leading macroeconometric model for the German economy.

[4]Strictly speaking, Gonzalo and Granger (1995) have suggested a permanent-transitory decomposition which is, in the understanding of this paper, a variant of trend-cycle measurement amongst a wide range of methods including statistical filters.

of segmentation, the macroeconomic factors influencing the trend in dwellings construction are derived in a less rigorous way. As shown in the second part, however, a cointegration analysis provides evidence which supports the claim that residential investment, population, per-capita real national income and real house prices form a long-run equilibrium relationship.

2.1 Some theoretical considerations

In contrast with the other private uses of GDP such as private consumption, business fixed investment and changes in inventories, macroeconomic theory has not established a leading hypothesis underpinning the econometric modelling of residential investment. This shortcoming might, in part, be explained by the ambiguous character of housing as a consumption and investment good.

From a utility perspective, private households consume the flow of services generated by a dwelling which they either own or rent. Whereas many consumption goods are exclusively used at the time of their purchase, dwellings possess an extremely long service life, suggesting that residential construction is a rather weak proxy for the use of housing services. Assuming that the services generated by an asset are proportionate to its value, this theory can be applied to the housing stock (net of depreciation), suggesting a positive correlation with income and wealth and a negative dependence on user costs. However, the inference for residential investment need not be straightforward, as comparatively long adjustment lags, non-negligible transaction costs, credit constraints and binding supply restrictions such as insufficient designation of building land and high market segmentation are factors to be considered.

While some side conditions could generally be integrated in a model of housing demand, it is the heterogeneity of the dwellings market which makes it difficult to recover hypotheses derived from consumption theory in aggregate data. Regional segmentation is one aspect worth mentioning in this regard. Perhaps equally important, in particular when residential construction is of primary concern, is the fact that the supply of housing is divided into owner-occupied and tenant-occupied dwellings.[5] Under real circumstances, the household's decision to build a house for its own use follows substantively different rules compared with an investor's decision to build, say, an apartment block with a number of rental units. Arbitrage mechanisms between house prices and rents are supposed to be present but work rather slowly owing to significant frictions. These include government interventions promoting either form of activity in residential construction.

Given the numerous difficulties in transferring the implications of a rigorous preference-based model of housing demand to a specification valid for a macroeconometric modelling exercise, let us opt for a more modest approach to discover the economic fundamentals explaining the trend in residential investment. In this context, the basic observation is that real dwellings construction has been increasing moderately over time, suggesting that,

[5]In 2003, 43 percent of private households lived in their own house or apartment. The ownership ratio had increased during the 1990s, with the ratio being significantly higher in western Germany. See Statistisches Bundesamt (2006) for details.

apart from cyclical fluctuations, there has been an ongoing capital formation process, given that the depreciation rate of residential buildings has remained more or less unchanged. The rising demand for housing services might be explained mainly by demographic factors and mounting economic prosperity.

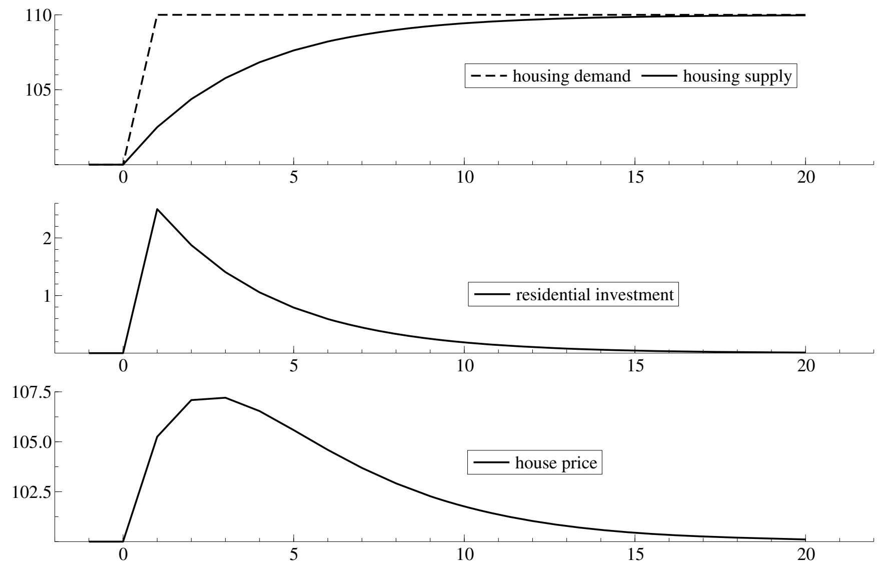
The diverse trends in the size and structure of population including ageing have a complex impact on housing demand. It seems *a priori* unclear whether the number of residents or the number of households is the better proxy in this respect. The number of households is directly connected with the number of occupied housing units but are neither all units equivalent in terms of investment costs nor does every household require the same unit. The living space increases with household size, albeit not in a one-to-one relationship owing to synergies in spatial use. While the square meters reserved for each household member and the creation of an own household are choices which follow economic principles based on preferences (relative to other consumption goods) and the budget constraint, housing is understood as a fundamental need of each individual regardless of her specific income position, justifying ascribing to demographic developments measured in terms of the population size a primary impact on the demand for housing.

Investments in new buildings and home improvements are expected as a consequence of rising economic prosperity. This is not only because individual households are then generally faced with laxer budget constraints and housing services tend to make up a larger share in their consumption bundles owing to the nature of housing as a superior good. Another reason is that the more prosperous a society is, the wider is generally the government's room for manoeuvre to assist poor households in terms of accommodation allowances and to provide enough subsidized public housing. Apart from demographic trends, the demand for housing is therefore assumed to be positively affected by the well-being of the society, which is approximated by a measure of economy-wide real income. In this context, an appropriate statistical concept is the gross national income adjusted for price and terms-of-trade effects.

Hypothesizing population and real national income as separate factors of housing demand may be disadvantageous in econometric settings because population and income are related to each other by a pure scale effect. While maintaining the primary importance of population, the scale effect in real national income can be removed by considering per-capita figures. As a consequence, the sensitivity of housing demand to changes in (absolute) income is reflected by the sensitivity to population provided that per-capita income remains unaltered. In this structure, the impact of population on housing demand is the amalgamation of the income effect and an additional (thereof independent) influence of demographic factors.[6] When demographic shifts are less than proportionately mirrored in the absolute level of income, housing demand is affected by the population effect net of a dampening impact which is related to the fact that per-capita income is reduced. It is assumed that the latter does not completely outweigh the former owing to effective means

[6] Under the simplifying assumption that the income elasticity of housing demand is unaffected by the level of income, the elasticity attributed to population is the sum of the income elasticity and the own elasticity of population.

Figure 1: Stylized time profile of housing market movements



of social policy assisting households which require additional space but face tight budget constraints.

The demand for housing is supposed to be a downward-sloping function of the real house price, where population and real national income per capita are regarded as location parameters shifting the function in a permanent manner. As determinants of a nonstationary nature, in a long-term perspective, they tend to dominate any other influences stemming from more cyclical factors such as user costs, lending conditions and credit availability. In formal terms, the demand for housing can be written as

$$H^D = H^D(\mathsf{PHR}; \mathsf{POP}, \mathsf{NIR}/\mathsf{POP}, X) \quad \text{with} \quad H_1^D < 0, H_2^D > 0, H_3^D > 0, \tag{1}$$

where PHR is the real house price, POP population and NIR gross national income adjusted for price and terms-of-trade effects. The vector X comprises the additional (short-term) impact factors; H_1^D, H_2^D and H_3^D denote the partial derivatives with respect to the argument and the first two location parameters of the function.

The supply of housing is assumed to be proportionate to the existing residential stock which evolves over time according to the accumulation identity

$$\Delta H^S = \mathsf{RI} - \delta H^S \tag{2}$$

where Δ is the difference operator, RI (gross) real residential investment and δ the depreciation rate, which is assumed to be constant over time. The stock of dwellings is expected to react rather slowly to variations in housing demand. This is due not only to the time it takes to construct buildings but also to many other factors holding up residential investments. These include the time households need to prepare and take the decision as well as delays owing to administrative aspects such as the designation of new building land and the grant of building permits.

The comparatively sluggish reaction of residential investment implies that, in the housing market equilibrium, house prices move as a result of tensions between demand and supply. As illustrated in Figure 1, an (unexpected) increase in housing demand which is triggered by, say, positive net migration immediately lifts house prices. So, the delayed adjustment of dwellings supply through positive (net) residential investment takes place amid elevated house prices. The reverse argument can be made for a negative demand shock.

Under real circumstances, the equilibrating tendencies may be quite persistent and are, thus, perceived as a medium-term to longer-term phenomenon. Apart from the slow pace of adjustment due to administrative obstacles and construction time, an ongoing price movement may automatically be reinforced by expectation mechanisms and speculation, occasionally leading to price bubbles in extreme cases. Rising house prices not only fuel residential construction to the extent that supply ultimately equilibrates demand but often also tend to induce over-investment owing to the expectation of high returns. By contrast, private households and investors abstain from constructing new dwellings when house prices are on a downward trend because expected capital losses reduce expected returns, making alternative uses of their funds more profitable.

Regardless of whether or not the reinforcing speculative component is present, the assumed sluggish adjustment of housing supply to changes in demand suggests a long-run positive comovement between residential investment and real house prices, with population and real income per capita playing a role as shifters of housing demand. Solved for residential investment, the long-run equilibrium relationship may be written as

$$\mathsf{RI}^* = f(\mathsf{POP}^*, \mathsf{NIR}^*/\mathsf{POP}^*, \mathsf{PHR}^*) \quad \text{with} \quad f_1 \geq 0, f_2 \geq 0, f_3 \geq 0, \tag{3}$$

where the superscript "$*$" indicates that the figures are equilibrium values and f_1, f_2 and f_3 denote the partial derivatives with respect to the arguments.

In the remainder of this paper, the pure population effect is referred to as the quantitative aspect of housing demand while per-capita real income is named as the qualitative component. As regards the measurement of house prices in real terms, it is worth mentioning that (nominal) house prices are deflated by the aggregate price index of all domestically produced goods except dwellings. This choice follows the idea of evaluating the price of new dwellings *vis-à-vis* the prices of any alternative uses in consumption and investment.

2.2 Cointegration analysis

Unit root tests provide evidence (see Appendix A) that the quarterly time series of residential investment, population, real national income per capita and real house prices, all series transformed into natural logarithms, are integrated of order 1, abbreviated by I(1),[7]

[7] In general, the order of integration d is henceforth denoted by I(d).

in the sample between the first quarter of 1975 and the fourth quarter of 2009. In the full sample, a statistical break in the first quarter of 1991 has to be taken into account, as the territorial basis changed at that date. For comparison, the overwhelming part of the empirical analysis is replicated in the west German subsample ending in the fourth quarter of 1991.[8] The existence of a long-run economic relationship of the form (3) can therefore be examined using cointegration analysis on the basis of a log-linearized version. In particular, it is tested whether cointegration can be established between the four series under consideration. More precisely, data should support the existence of exactly one cointegrating relation, with the coefficients taking the theoretically hypothesized signs.

The variables of interest are jointly modelled by a vector autoregression (VAR) of lag order p including deterministic regressors. If the full sample is considered, the system comprises a constant, a linear trend and a step dummy variable $\mathsf{S(91{:}1)}$, which is unity since the first quarter of 1991 and otherwise zero. The latter variable is absent when the analysis is restricted to the west German subsample, suggesting that the standard likelihood ratio (LR) cointegration tests proposed by Johansen (1991) can be applied. In the full sample, however, it is necessary to account for the existence of a structural break in the first quarter of 1991. An appropriate handling of this element in the set framework is to use the cointegration test proposed by Saikkonen and Lütkepohl (2000), henceforth abbreviated by SL. The basic idea of this approach is to remove the deterministic components prior to testing for cointegration using a standard LR test *à la* Johansen (1991). The linear trend is supposed to be absent in the cointegrating relation, as (3) does not suggest any role for it in the long-run equilibrium relationship. This hypothesis in turn implies that, in a K-dimensional system, the cointegration rank r is $K-1$ at maximum, given that the involved time series are trending over time.

When cointegration is established, the resulting cointegrating relation represents (3) in log-linearized form. It may be written as

$$\mathsf{ri}_t - b_1\mathsf{pop}_t - b_2\mathsf{inc}_t - b_3\mathsf{phr}_t + c = \mathsf{ec}_t \quad \sim \quad \mathrm{I}(0), \tag{4}$$

where ri_t, pop_t, inc_t and phr_t are the logs of RI_t, POP_t, $\mathsf{NI}^*_t/\mathsf{POP}^*_t$ and PHR^*_t respectively. Theory suggests b_1, b_2, b_3 to be positive or nil. The error correction term or long-run residual series is denoted by ec_t, with the constant c ensuring that this series possesses a zero mean.

Cointegration testing follows the specific-to-general rule, implementing the logic that there is no role for any further I(1) variable in a single cointegrating relation if cointegration has already been established between some I(1) variables (Lütkepohl, 2007). Given the primary importance of population for residential investment, this means that the smallest system is two-dimensional, modelling only the quantitative element of housing demand without a persistent relative price effect. If cointegration between residential investment and population cannot be established, there are two alternatives for a three-dimensional system. The first is $\{\mathsf{ri}, \mathsf{pop}, \mathsf{inc}\}$, meaning that the trend in residential investment is

[8] In this subsample, the data for the year 1991 refers to western Germany while it refers to Germany as a whole in the full sample.

Table 1: Cointegration tests

lag order	$p = 2$ (HQ, SC)			$p = 5$ (AIC)		
null hypothesis	$r = 0$	r ≤ 1	$r \leq 2$	$r = 0$	$r \leq 1$	$r \leq 2$
(a) Germany (1975 – 2009)						
{ri, pop}	7.47 [0.134]			5.22 [0.311]		
{ri, pop, inc}	16.23 [0.195]	5.50 [0.282]		13.18 [0.394]	4.99 [0.338]	
{ri, pop, phr}	27.02 [0.006]	2.55 [0.706]		26.09 [0.009]	5.00 [0.337]	
{ri, pop, inc, phr}	39.99 [0.016]	12.96 [0.411]	3.94 [0.478]	32.89 [0.030]	16.12 [0.200]	5.06 [0.329]
(b) western Germany (1975 – 1991)						
{ri, pop}	11.04 [0.213]			11.17 [0.204]		
{ri, pop, inc}	22.19 [0.297]	6.72 [0.616]		21.70 [0.326]	10.20 [0.270]	
{ri, pop, phr}	21.34 [0.347]	8.75 [0.396]		21.33 [0.348]	8.87 [0.385]	
{ri, pop, inc, phr}	52.80 [0.015]	26.96 [0.105]	10.67 [0.237]	54.71 [0.009]	26.98 [0.105]	10.39 [0.256]

Cointegration tests are performed in VARs of lag order p. The full-sample specifications include a step dummy and a (contemporaneous) impulse dummy capturing the statistical break in the first quarter of 1991. The p-values of the LR trace tests taken from the software JMulti 4 are reported in brackets.

represented by the quantitative and the qualitative determinants of housing demand, and the second is {ri, pop, phr}, implying the existence of the relative price effect in a long-run equilibrium where housing demand is restricted to the quantitative aspect only. The full model contains all variables; thus, ultimately, an attempt is made to establish cointegration in a four-dimensional system.

Table 1, Panel (a), shows the results of the SL cointegration tests carried out in the whole sample for the different systems under review. They are performed in VARs of lag orders 2 and 5, with the parsimonious lag length representing the choices of the consistent information criteria HQ and SC while the long order results from AIC selection (Lütkepohl, 2005).[9] It is evident that the trend in residential investment cannot be modelled by relying solely on the factors of housing demand. It is necessary to include real house prices in order to establish cointegration. As this variable is understood as a proxy for the tensions between supply and demand, its presence in the cointegrating relation implies that the forces equilibrating the quantities on the housing market work rather slowly. Hence, price movements in either direction are persistent enough to be empirically modelled by a unit

[9] All cases are chosen out of a set of VARs whose lag orders take the integers from 1 to 10.

Table 2: Cointegrating relations

sample	*Germany (1975 – 2009)*		*western Germany (1975 – 1991)*	
method	OLS	DOLS	OLS	DOLS
ri_t	1	1	1	1
pop_t	-1.56 (0.06)	-1.61 (0.06)	-2.58 (0.43)	-2.37 (0.51)
inc_t	-0.26 (0.08)	-0.26 (0.08)	-0.38 (0.05)	-0.38 (0.08)
phr_t	-0.76 (0.13)	-0.83 (0.13)	-0.64 (0.08)	-0.66 (0.09)
const	-7.4 (0.7)	-8.0 (0.7)	11.3 (1.7)	10.5 (2.1)
adj. R^2	0.92	0.92	0.77	0.77
DW	0.28	0.25	0.84	0.81

Standard errors are reported in parentheses. Auxiliary regressors in DOLS estimates are omitted.

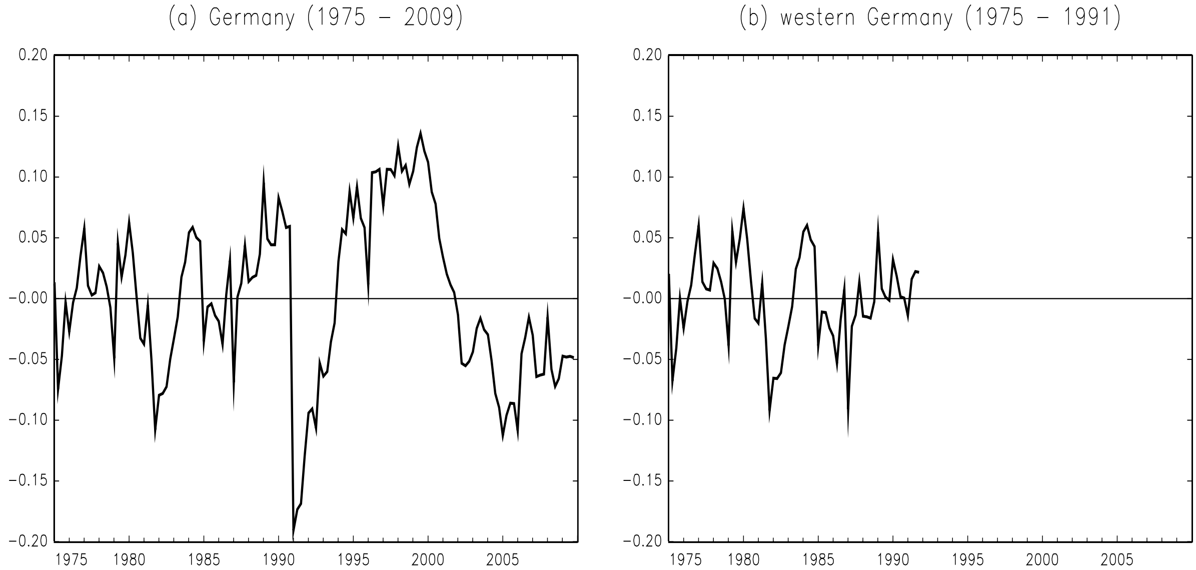
root process. Furthermore, the cointegration analysis for the full sample indicates that the qualitative aspect of housing demand is rather weak, as cointegration is already present in the three-dimensional system $\{\mathsf{ri}, \mathsf{pop}, \mathsf{phr}\}$.

By contrast, the quality aspect plays a crucial role in explaining the trend in housing demand when only the west German subsample is considered. As displayed in Table 1, Panel (b), Johansen's (1991) standard LR test procedure finds evidence for the existence of exactly one cointegrating relation between the four variables at conventional significance levels, whereas cointegration is absent in all two-dimensional or three-dimensional subsystems under consideration. However, the results from the reduced sample should be interpreted with caution because, with only 68 observations, cointegration rank tests are less reliable than the full-sample analysis, in particular when the systems are of higher dimension.

The long-run relationship between residential investment, the trend determinants of housing demand and real house prices is established by estimating the static regression choosing residential investment as the left-hand side variable. Apart from using ordinary least squares (OLS), the dynamic OLS (DOLS) approach is additionally applied in order to come closer to estimation efficiency (Saikkonen, 1991). Starting from a model including the first differences of the (exogenous) regressors together with their first two leads and lags, specification search using standard information criteria suggests maintaining the contemporaneous differences only.[10]

[10] A maximum likelihood estimation of the full system, however, yields less reliable results because the number of coefficients to be estimated is high compared with the number of observations available, in particular as far as the west German subsample is concerned.

Figure 2: Long-run residuals of residential investment



Regardless of whether the full sample or the west German subsample is considered, the coefficients of the cointegrating relation show the expected signs. The estimates do not differ much depending on whether the OLS or the DOLS approach is used. While cointegration testing in the full sample has not suggested that per-capita real income is necessarily present in the long-run equilibrium relationship, the estimates of the cointegrating vector point to a rather small but not statistically negligible impact. The essential role of per-capita real income in the cointegrating relation for the west German era is confirmed by a higher coefficient (in absolute value).

The regression results also suggest that the impact of population on residential investment is more than a pure scale effect, as the estimated coefficient clearly exceeds unity (in absolute value). Hence, the trend in residential investment seems to be quite sensitive to demographic developments. While a population elasticity which exceeds unity may be difficult to be theoretically reconciled with households' budget constraints in an economy possessing a high fertility rate, it may be sustainable in the German context where, apart from some pronounced waves of immigration, the population has been more or less stable for the last 35 years and the qualitative aspect of housing demand turns out to be of rather limited magnitude.

The significance of real house prices in the cointegrating relation confirms the view that persistent tensions between housing demand and supply are reflected in the price of new dwellings, with the sign of the coefficient signalling that supply tends to adjust slowly to variations in demand. Here, too, the empirical results are consistent with the suggestions developed in the theoretical part of the paper.

As displayed in Figure 2, what is striking in the long-run residual series (taken from the DOLS estimates) is that the degree of persistence changed substantially at reunification. Whereas fluctuations of a standard business cycle length are apparent in the west German era, the residual series thereafter describes a pronounced cycle lasting from 1991 to 2005 followed by a mild recovery. The prolonged cycle duration is also reflected by the fact that serial correlation is distinctly higher in the full-sample residual series compared with its subsample counterpart, which is reflected by a substantially lower Durbin-Watson (DW) statistic.

In sum, the cointegration analysis has provided two main results. First, the trend in residential investment cannot be explained solely by fundamental factors of housing demand such as population growth and permanent gains in per-capita real national income. To establish cointegration, it is necessary to additionally consider real house price developments as a measure of persistent tensions between housing demand and supply. Second, while this pattern is generally maintained over the whole sample under investigation, reunification led to a change in the degree of persistence of the error correction term.

3 Cycle features of residential investment

The analysis of this section is based on the observation that the residual series of the cointegrating relation established between residential investment, population and real national income per capita and real house prices has changed the degree of persistence in the course of reunification. This fact matters for the measurement of activity cycles in the housing market.

The in-depth study of cyclical features in residential investment generally uses the framework of a vector error correction model (VECM). However, central results can be derived from a specification which is substantively simplified owing to the observation that population, per-capita real income and real house prices are weakly exogenous. The preferred model structure also makes it possible to form the trend-cycle decomposition suggested by Gonzalo and Granger (1995) with observable variables and the estimated cointegrating vector only. In particular, the Gonzalo-Granger (GG) cycle component of residential investment is defined by the residual series of the cointegrating relation. The final part of this section shows that the GG cycle deviates substantively from the results of standard univariate trend-cycle filters, which gives rise to a discussion on the pitfalls of simple filter techniques in a situation where the properties of the time series to be filtered change.

3.1 Adjustment to the housing market equilibrium

The study of adjustment processes towards the long-run housing market equilibrium requires a complete econometric specification of the time series under review. Accounting for cointegration, the VAR(p) representing the data generating process of the four variables stacked in $y_t = (\mathsf{ri}_t, \mathsf{pop}_t, \mathsf{inc}_t, \mathsf{phr}_t)'$ can be rewritten as the following VECM

$$\Delta y_t = \alpha \mathsf{ec}_{t-1} + \sum_{j=1}^{p-1} \Gamma_j \Delta y_{t-j} + \Phi D_t + \varepsilon_t, \tag{5}$$

where α is the four-dimensional vector of adjustment parameters correcting non-zero realizations of the (scalar) long-run residual series. It is taken from (4), which is compactly written as $\mathsf{ec}_t = \beta' y_t + c$ with $\beta = (1, -b_1, -b_2, -b_3)'$. In addition, $\Gamma_1, ..., \Gamma_{p-1}$ are (4×4) parameter matrices, Φ is the parameter matrix attached to the deterministic components collected by D_t which includes const and impulse dummy variables related to S(91:1). The residual term ε_t is assumed to follow a Gaussian vector white noise process with zero mean and the variance-covariance matrix Ω.

A variable may contribute to the convergence of the system towards its long-run equilibrium if the coefficient attached to the error correction term in the respective system equation is significantly different from zero. A time series which does not fulfill this requirement is said to be weakly exogenous (with respect to the cointegrating space). A LR test for weak exogeneity, provided that cointegrating vectors are given or estimated, is suggested by Johansen (1996). For a variable to contribute to the error correction process, however, it is sufficient that the adjustment parameter and the coefficient attached to this variable in the cointegrating vector possess opposite signs. The speed of the adjustment process is reflected in the magnitude of the parameter in absolute value.

Table 3 reports the estimates of the adjustment parameter vector and the results of weak exogeneity tests based on the super-consistent estimates of the cointegrating vector. While the results do not vary with the estimation methods, the properties of the model alter considerably depending on whether the full sample or the west German subsample is considered. In the full sample, the error correction term only loads on $\Delta\mathsf{ri}_t$, while the adjustment parameters in the other system equations are not significantly different from zero at conventional significance levels. Hence, it comes as no surprise that the null hypothesis of weak exogeneity of $\{\mathsf{pop}_t, \mathsf{inc}_t, \mathsf{phr}_t\}$ is not rejected by a formal test.

If the analysis is restricted to the west German subsample, results change qualitatively, as deviations from the long-run equilibrium relationship are, apart from residential investment, also corrected via real house prices. The equilibrium-correcting force of the price variable is less strong (in absolute value). However, the adjustment parameter in the system equation of $\Delta\mathsf{phr}_t$ is significantly different from zero and weak exogeneity of $\{\mathsf{pop}_t, \mathsf{inc}_t, \mathsf{phr}_t\}$ is clearly rejected, too. By contrast, the joint hypothesis that population and per-capita income are weakly exogenous cannot be rejected by a statistical test. Thus, in the years between 1970 and 1991, residential construction and house prices helped "correct" disequilibria between housing demand and supply.

Since reunification the endogeneity of house prices has been lost and the speed of the error correction mechanism via residential investment has slowed down considerably. The adjustment parameter attached to $\Delta\mathsf{ri}_t$ is about four times lower in the full-sample estimate than in the estimate on the basis of the west German subsample. More precise evidence on this issue can be obtained by specifying a conditional error correction model in residential

Table 3: Adjustment parameter estimates and weak exogeneity tests

sample	*Germany (1975 – 2009)*		*western Germany (1975 – 1991)*	
method	OLS	DOLS	OLS	DOLS
(a) adjustment parameters				
$\Delta\mathsf{ri}_t$	-0.108 (0.039)	-0.109 (0.039)	-0.390 (0.121)	-0.380 (0.122)
$\Delta\mathsf{pop}_t$	0.002 (0.002)	0.002 (0.002)	0.001 (0.002)	0.001 (0.002)
$\Delta\mathsf{inc}_t$	-0.002 (0.013)	0.001 (0.013)	-0.065 (0.040)	-0.059 (0.040)
$\Delta\mathsf{phr}_t$	0.003 (0.008)	0.004 (0.008)	0.061 (0.026)	0.063 (0.026)
(b) weak exogeneity				
$\{\mathsf{pop}_t, \mathsf{inc}_t, \mathsf{phr}_t\}$	2.61 [0.456]	2.73 [0.436]	8.35 [0.039]	8.35 [0.039]
$\{\mathsf{pop}_t, \mathsf{inc}_t\}$	2.46 [0.292]	2.42 [0.298]	2.83 [0.243]	2.50 [0.287]

Panel (a) presents the estimates of the vector α in (5); standard errors are reported in parentheses. Panel (b) reports the LR test statistics of the weak exogeneity hypothesis. The statistics are asymptotically χ^2 distributed, with the degrees of freedom equalling the number of variables tested to be weakly exogenous. The p-values of these tests are reported in brackets. In both cases, results are subject to prior OLS/DOLS estimation of the cointegrating vector.

investment for the full sample, letting the adjustment parameter switch at the date of reunification. According to Johansen (1992), this setup provides efficient estimates of the parameters of interest, as population, per-capita real national income and real house prices can be assumed to be weakly exogenous.

Without loss of generality, the (full) VECM in (5), rewritten in the form

$$\begin{bmatrix} \Delta y_{1t} \\ \Delta y_{2t} \end{bmatrix} = \begin{bmatrix} \alpha_1 \\ \alpha_2 \end{bmatrix} \mathsf{ec}_{t-1} + \sum_{j=1}^{p-1} \begin{bmatrix} \Gamma_{1j} \\ \Gamma_{2j} \end{bmatrix} \Delta y_{t-j} + \begin{bmatrix} \Phi_{1j} \\ \Phi_{2j} \end{bmatrix} D_t + \begin{bmatrix} \varepsilon_{1t} \\ \varepsilon_{2t} \end{bmatrix}, \quad \Omega = \begin{bmatrix} \Omega_{11} & \Omega_{12} \\ \Omega_{21} & \Omega_{22} \end{bmatrix}, \qquad (5')$$

can be decomposed into the conditional model

$$\Delta y_{1t} = \alpha^c \mathsf{ec}_{t-1} + \omega \Delta y_{2t} + \sum_{j=1}^{p-1} \Gamma_j^c \Delta y_{t-j} + \Phi^c D_t + \varepsilon_t^c \qquad (6)$$

and the marginal model

$$\Delta y_{2t} = \alpha_2 \mathsf{ec}_{t-1} + \sum_{j=1}^{p-1} \Gamma_{2j} \Delta y_{t-j} + \Phi_2 D_t + \varepsilon_{2t}, \qquad (7)$$

where $\omega = \Omega_{12}\Omega_{22}^{-1}$, $\alpha^c = \alpha_1 - \omega\alpha_2$, $\Gamma_j^c = \Gamma_{1j} - \omega\Gamma_{2j}$, $\Phi^c = \Phi_1 - \omega\Phi_2$, and $\varepsilon_t^c = \varepsilon_{1t} - \omega\varepsilon_{2t}$ is a white noise process with the variance-covariance matrix $\Omega_{11}^c = \Omega_{11} - \omega\Omega_{21}$. Under weak exogeneity of the variables stacked in y_{2t} (*i.e.* $\alpha_2 = 0$), it is $\alpha^c = \alpha_1$ and statistical inference on α_1 can be based on (6) only.

Following the results of the weak exogeneity tests, let $y_{1t} = \mathsf{ri}_t$ and $y_{2t} = (\,\mathsf{pop}_t, \mathsf{inc}_t, \mathsf{phr}_t\,)'$. In addition, α_1 is allowed to switch at reunification. Assuming lag order $p = 1$ but subsequently dropping nonsignificant coefficients in ω, Γ_1^c and Φ^c, an OLS regression of this specification using the DOLS estimate of the cointegrating relation (4) yields

$$\begin{aligned} \Delta\mathsf{ri}_t = &-\underset{(0.075)}{0.250}\,\mathsf{ec}_{t-1} + \underset{(0.081)}{0.184}\,[\,\mathsf{ec}_{t-1}\,\mathsf{S(91{:}1)}\,] + \underset{(1.65)}{4.32}\Delta\mathsf{pop}_t + \underset{(0.18)}{0.86}\Delta\mathsf{inc}_t + \underset{(0.26)}{0.52}\Delta\mathsf{phr}_t - \\ &- \underset{(0.38)}{0.75}\Delta\mathsf{S(91{:}1)} - \underset{(0.006)}{0.029}\mathsf{ice1}_t - \underset{(0.007)}{0.013}\mathsf{ice4}_t + \underset{(0.006)}{0.021}\mathsf{ice1}_{t-1} + \underset{(0.007)}{0.012}\mathsf{ice4}_{t-2} + \varepsilon_t^c, \qquad (8) \\ &\mathrm{R}^2 = 0.56, \qquad \mathrm{DW} = 2.01, \end{aligned}$$

where $\mathsf{ice1}_t$ and $\mathsf{ice4}_t$ measure the number of days with a maximum temperature below freezing in the first and the fourth quarter respectively as a percentage of the seasonal mean and are zero in the remaining quarters.

The estimate highlights the significant drop in the adjustment parameter in absolute value, from 0.250 in the west German era to 0.066 afterwards.[11] The built-in forces for equilibrium correction in the housing market via residential investment have therefore weakened substantively since reunification. In addition, residential investment is instantaneously affected by changes in population and per-capita real national income. The short-term elasticity of residential investment to demographic variations is particularly high.

Construction activity is hampered (favoured) by an unusually severe (mild) winter.[12] The weather effect is revealed to be roughly twice as strong when it appears in the first quarter rather than the fourth quarter. The significantly positive coefficients attached to $\mathsf{ice1}_{t-1}$ and $\mathsf{ice4}_{t-2}$ point to the existence of countereffects in the following spring. The estimates of the coefficients are plausible also in their (absolute) magnitude, as the positive lagged effects more or less compensate for the initial negative effects.

3.2 A trend-cycle decomposition of residential investment

The VECM in (5) implies a permanent-transitory decomposition of the Gonzalo and Granger (1995) type. It is interpreted as an approach to trend-cycle measurement where

[11] Although the resulting estimate of the adjustment parameter is rather low for the pan-German era, it turns out to be statistically significant from zero, as a Wald-type coefficient test for equal but reversely signed coefficients attached to ec_{t-1} and $\mathsf{ec}_{t-1}\mathsf{S(91{:}1)}$ yields the statistic 4.45, which suggests that the null hypothesis is rejected at the 5% level.

[12] Recall that the analysis considers residential investment in seasonally and working-day adjusted form, implying that the dampening effect of a "usual" winter on residential construction is removed by seasonal adjustment.

the identification follows the rule that, in the long-run (*i.e.* for horizons approaching infinity), conditional predictions of y_t are exclusively affected by shocks to trend factors while shocks to cycle components only possess a temporary impact.

For the formal description, let us denote by $\alpha_\perp$ and $\beta_\perp$ the orthogonal complements of α and β respectively, satisfying $\alpha'_\perp\alpha = 0$ and $\beta'_\perp\beta = 0$. The GG trend-cycle decomposition is then given by

$$y_t = y_0 + y_0^b \mathsf{S(91{:}1)} + A\tau_t + Bc_t \quad \text{with} \quad \tau_t = \alpha'_\perp y_t \quad \text{and} \quad c_t = \beta' y_t, \tag{9}$$

where y_0 and y_0^b are some initial values, the three-dimensional vector process τ_t represents the trend factors and c_t is the time series of the single cyclical factor in this system. The (4×3) matrix $A = \beta_\perp(\alpha'_\perp\beta_\perp)^{-1}$ collects the loadings attached to the trend factors while the loadings attached to the cyclical factor are stacked in $B = \alpha(\beta'\alpha)^{-1}$, which is a four-dimensional vector in the case of a single cointegrating relation.

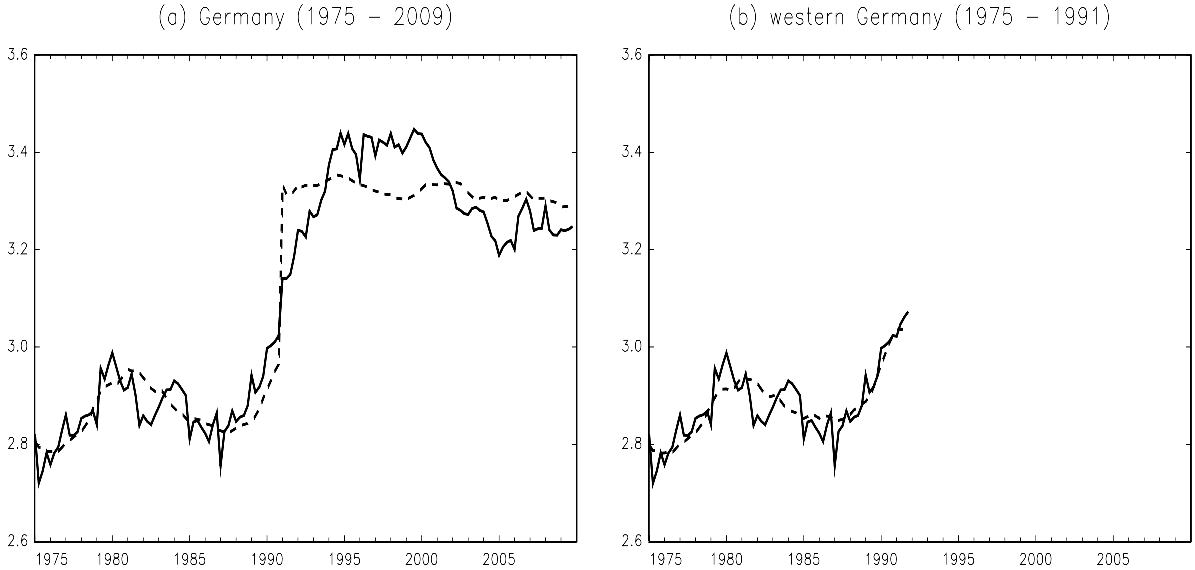
Under weak exogeneity of pop_t, inc_t and phr_t, *i.e.* $\alpha = (a_1, 0, 0, 0)'$ with $a_1 < 0$, (9) simplifies substantively. First, the weakly exogenous variables directly represent the three independent stochastic trend factors driving residential investment because, in this case, $\alpha_\perp = [o : I_3]'$ where I_3 denotes an identity matrix of order 3 and o is a three-dimensional vector of zeros. The trend component of residential investment is the weighted average of the weakly exogenous variables where the weighting scheme is given by the coefficients in the cointegrating vector attached to them with reverse sign. Second, the cyclical factor only loads on ri_t because $B = (1, 0, 0, 0)'$. Hence, the cycle component of residential investment is represented by the error correction term. In sum, (9) can be written as

$$\begin{bmatrix} \mathsf{ri}_t \\ \mathsf{pop}_t \\ \mathsf{inc}_t \\ \mathsf{hpr}_t \end{bmatrix} = \begin{bmatrix} c \\ 0 \\ 0 \\ 0 \end{bmatrix} + \begin{bmatrix} b_1 & b_2 & b_3 \\ 1 & 0 & 0 \\ 0 & 1 & 0 \\ 0 & 0 & 1 \end{bmatrix} \begin{bmatrix} \mathsf{pop}_t \\ \mathsf{inc}_t \\ \mathsf{hpr}_t \end{bmatrix} + \begin{bmatrix} \mathsf{ec}_t \\ 0 \\ 0 \\ 0 \end{bmatrix}. \tag{9'}$$

In this framework, the GG trend-cycle decomposition is informative for residential investment only, while population, per-capita real national income and real house prices do not possess a meaningful cycle component. It is worth stressing that the latter set of variables may nonetheless include transitory dynamics; however, these fluctuations lie in the space spanned by $\alpha_\perp$ and, thus, fail to possess the capacity to bring about equilibrium correction (Proietti, 1997). The fact that the speed of convergence via residential investment has slowed down since reunification does not affect the trend-cycle decomposition in a formal sense because neither $\alpha_\perp$ nor B depend on α in this special case. Instead, it is the time series structure of the error correction term which causes the increase in the degree of persistence characterizing the cycle component of residential investment since 1991.

Figure 3 shows residential investment *vis-à-vis* its GG trend component, resulting from the estimates in the full sample and the west German subsample. Overall, the trend is smooth but fluctuates visibly as a result of cycling behavior in fundamental factors. Most prominent in this respect are the waves of immigration into western Germany at the end of the 1980s as well as into Germany in the early 1990s. In the west German era, phases of

Figure 3: Gonzalo-Granger trend component of residential investment



Actual residential investment is displayed by the solid line and the GG trend components by dashed lines.

economic weakness as in the mid-1970s and in the early 1980s were associated with gaps in residential construction.

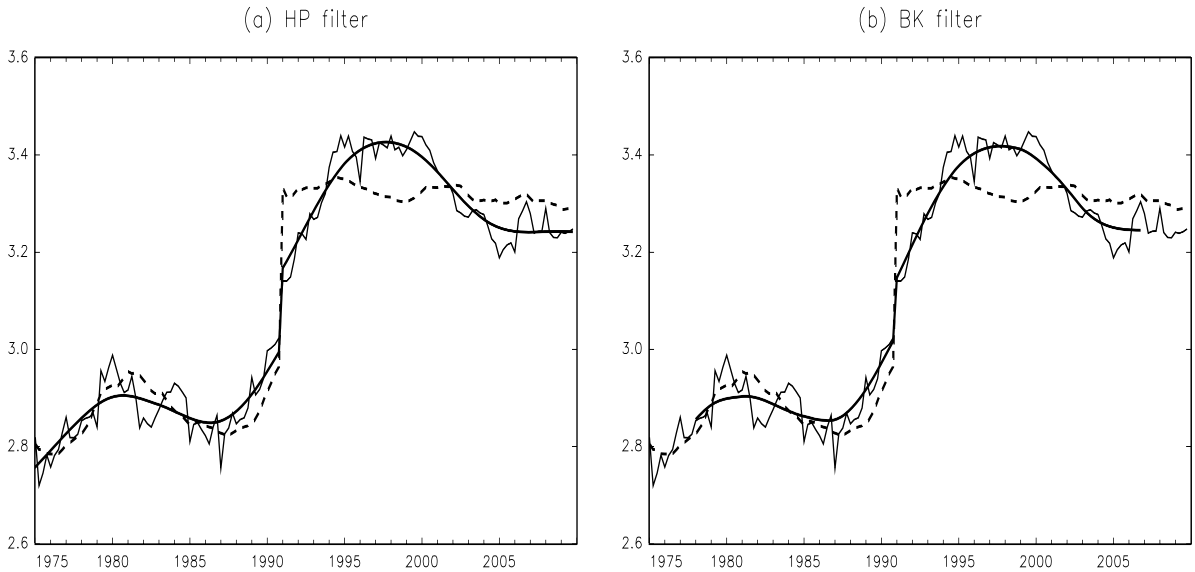
While departing from a more or less balanced cyclical position in 1990 despite rather strong trend growth just before that, the reunification-induced shift in the fundamentals of housing demand, in particular the increase in residents, required a considerably higher trend level of dwellings construction only half of which could be promptly served by existing capacities. Consequently, the construction sector expanded rapidly, lifting residential investment not only to its trend but far beyond it. The excessive formation of residential capital during the 1990s called for an adjustment which, in turn, materialized in the form of a long period of below-trend construction activity. This pattern prevailed throughout the past decade though the distance to the trend level has reduced significantly since 2005.

The GG trend of residential investment is virtually flat between 1991 and 2009. Relying on the econometric description in (9′) informed by the development of economic fundamentals, this observation is ascribed to anemic demand for housing, as the population stagnated virtually and real national income per capita evolved less dynamically. With an exception towards the end of the 1990s, real house prices also declined steadily in this period.

3.3 Pitfalls for univariate statistical filtering

The change in the cyclical features of residential investment makes it difficult for univariate statistical filter techniques to come out with a trend-cycle decomposition resembling the

Figure 4: Trend extraction with univariate statistical filters



The HP and BK trend components of residential investment are displayed by the thick solid line in the respective charts. Actual residential investment (thin solid line) and the GG trend component (dashed line) are shown as reference series.

one that was derived in the previous section. The reason is that univariate filters lack flexibility in the sense that they extract the trend component on the basis of a time-invariant statistical criterion. For instance, the Hodrick-Prescott (HP) filter imposes a constant variance ratio λ between trend innovations and shocks to the cycle, whereas the Baxter-King (BK) filter defines a fixed frequency limit below which oscillations are ascribed to the trend component.

Figure 4 shows the trend component of residential investment resulting from the HP filter with $\lambda = 1,600$ (as usual for quarterly data) and from the BK high-pass filter assigning to the trend those oscillations whose duration surpasses eight years. The reunification shift is incorporated in the HP procedure by including S(91:1) in the measurement equation of the underlying state-space model. The BK trend component results from applying the high-pass filter to the residuals of the OLS regression of residential investment on const and S(91:1).

While the outcomes of the statistical filtering techniques are virtually the same, they differ markedly from the GG trend component. In general, the HP and BK trends follow more closely the actual movements in residential investment, making the corresponding cyclical fluctuations less pronounced. This feature is most striking after reunification, as the boom-bust movement in dwellings construction is attributed to the HP and BK trends, whereas the GG trend is flat. In addition, the cointegration analysis shows that the economic fundamentals of housing demand would have implied a substantially bigger mean shift in residential investment than the one that actually happened at reunification. By contrast, the removal of the structural break incorporated in the HP filtering exercise brings about a mean shift in the resulting trend components which closely resembles the observed change in residential investment between the west German level of the fourth quarter of 1990 and the pan-German level of the first quarter of 1991.[13]

The marked discrepancy between the GG trend-cycle decomposition and the results of the univariate statistical filter techniques can be further highlighted by studying the time series characteristics of the corresponding cycle components. Owing to the strong similarity of the HP and BK results, the comparative analysis is restricted to the HP cycle.[14] Figure 5 plots the GG and HP cycle components of residential investment, which differ from the gap between the actual series and the respective trend components in that, in order to reduce erratic variations, the impact of unusual weather conditions in the winter term including the counter effects in the following spring are eliminated by a regression approach.[15] The two cycle components differ in terms of duration and amplitude. In addition, the time series characteristics of the HP cycle component seem to be stable over time while, in the GG cycle component, a switch is recognizable in the course of reunification.

In order to confirm the visual impression with statistical evidence, spectral analysis is used. The idea is to find significant differences in shape between the full-sample spectrum and the spectra of the west-German and pan-German subsamples respectively. As the full-sample spectrum serves as a reference, its point estimate is surrounded by 95 percent confidence bands which are constructed using the bootstrap procedure suggested by Franke and Härdle (1992). With this technique, kernel spectral estimates are bootstrapped by resampling from the periodogram of the data. The bootstrap is based on 2,000 replications.

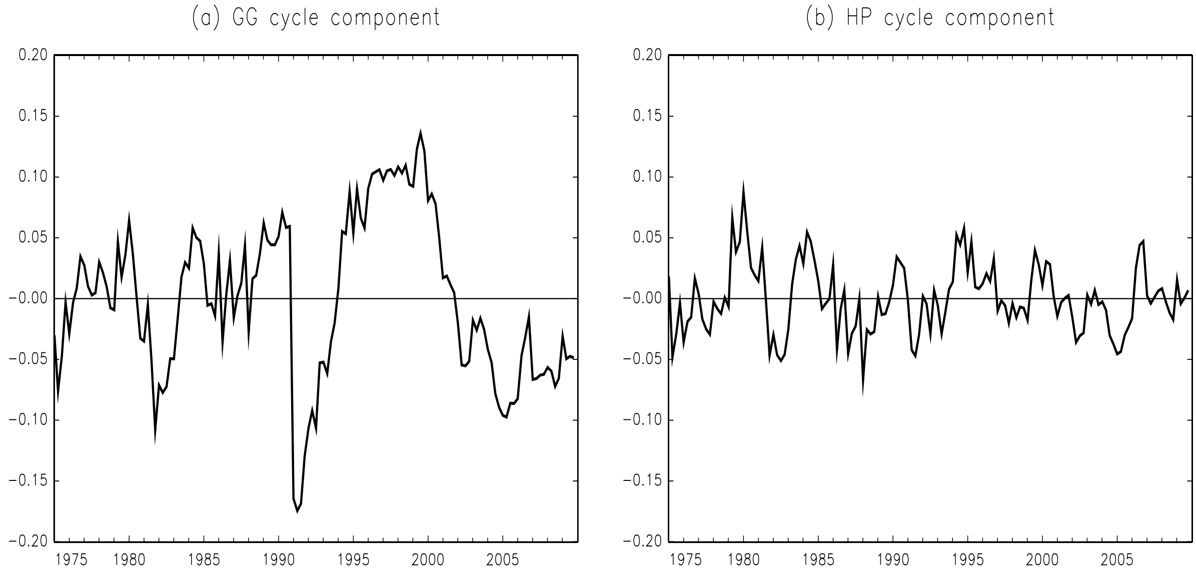
As shown in the upper panel of Figure 6, the full-sample spectrum of the GG cycle component peaks in a frequency range which is consistent with a cycle duration of about eight years. The observation that the peak in the spectrum of the west German subsample is located to the right while the pan-German counterpart is located to the left of the eight-years frequency implies that the cycle length of the dominating oscillation increased at reunification, confirming the visual impression taken from the time series plot. In addition, the spectra of the west German and pan-German subsamples also differ with respect to the height of peak. This observation means that the spectral mass attached to the dominating oscillation is much more concentrated in the GG cycle component after reunification than in the spectrum of the west German subsample, suggesting that fluctuations in the range of major business cycles (and beyond it) shape the cyclical behavior of residential investment after reunification to a higher degree than before. Considering the confidence bands of the full-sample estimates, the evidence is significant from a statistical point of view, as the peak in the spectrum of the pan-German subsample clearly surpasses the upper limit

[13] The mean shift in the BK trend is virtually equal to the mean shift in the actual series.

[14] The HP filter has been chosen for simplicity, namely because of the fact that, in contrast to the BK filter, it provides results for the full sample without any further treatment of data or techniques.

[15] The regressors are $\mathsf{ice1}_t$ and $\mathsf{ice4}_t$ including the countereffect in the forthcoming spring, which have been shown to possess a significant effect on the quarter-on-quarter changes in residential investment.

Figure 5: GG and HP cycle components of residential investment



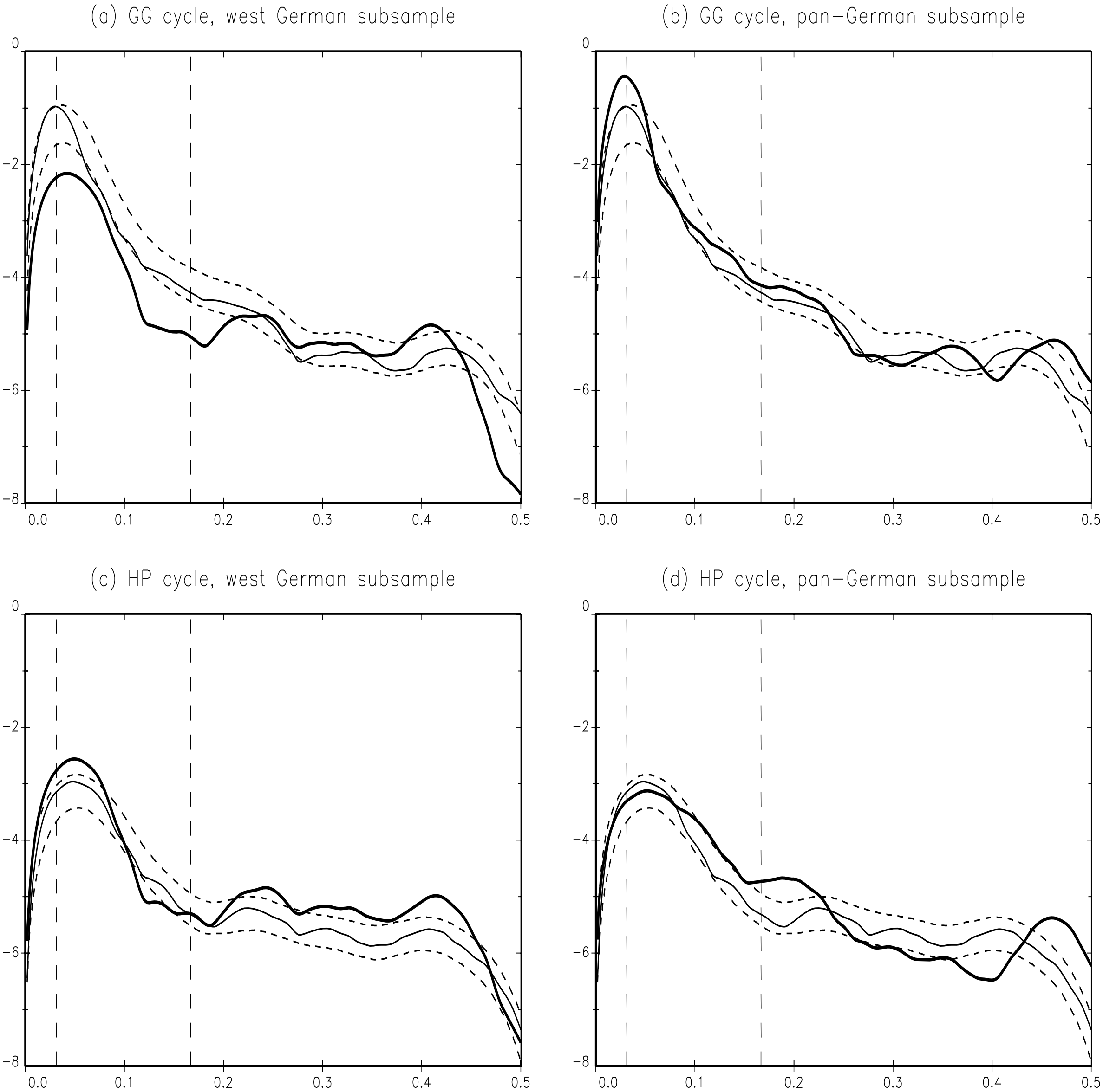
The cycle components are measured as percentage deviation from the trend.

of the confidence interval while the peak in the spectrum of the west-German subsample drops below the lower limit.

The frequency-domain analysis of the HP cycle component reveals a number of differences in comparison with the GG cycle component. First, the HP cycle spectrum peaks visibly within the range of business cycle frequencies. In particular, the duration of the oscillation to which most spectral mass is assigned is about five years. Second, the HP spectrum is generally flatter than the GG spectrum, suggesting that, in the HP cycle component, fluctuations at business cycle frequencies are less dominant relative to erratic variations. Third, the subsample spectra of HP cycle component do not exhibit striking differences in the placement and the height of the peak. In the frequency range around the peak, the spectrum of the pan-German subsample lies within the confidence interval of the full-sample spectrum while the spectrum of the west-German subsample slightly exceeds the upper limit. A substantive break in the cyclical properties of the HP cycle component cannot be concluded from this evidence.

In sum, the changing cyclical pattern in residential investment, which is a key result of the econometric analysis using a cointegrated VAR framework and which can be represented in terms of a GG trend-cycle decomposition, cannot be replicated by standard univariate filter techniques such as the HP and BK filter. Taking for granted the conclusions from the cointegration analysis, which incorporates a great deal of theoretical considerations, this can be seen as a deficiency and, thus, caution is advised when interpreting and using results which are derived from HP or BK filtered time series in this context.

Figure 6: Spectra of cycle components



The charts depict the spectrum of the subsample mentioned in the respective headline as a thick solid line. The point estimate of the full-sample spectrum is drawn by the thin solid line, with the limits of the 95 percent confidence interval represented in the dashed format. The abscissa scale is frequency divided by 2π. The dashed vertical lines limit the frequency band attributed to periodicities of two and eight years ("business cycle frequencies").

4 Conclusion

Real residential investment in Germany is modelled in a cointegrated vector autoregression with population, per-capita real national income and real house prices . Regarding the estimates of the cointegrating relation, demographic factors affect dwellings construction through more than a pure scale effect, while real income per capita turns out to have a small but statistically non-negligible impact. Tensions between housing demand and supply seem to be persistent, assigning real house prices a significant role in the long-run equilibrium relationship. The long-run residual series which is implied by the cointegrating vector is subject to a marked increase in the degree of persistence at reunification.

The specification and estimation of the full vector error correction model have given rise to the belief that, at reunification, the speed of equilibrium adjustment via residential investment slowed down and real house prices lost the capacity to contribute to the adjustment process. If the analysis is based on the full sample including west German data from 1970 to 1990 and pan-German data from 1991 onward, population, real income per capita and real house prices are weakly exogenous. This result implies that the full time series properties of residential investment can be described by a single-equation error correction model conditioned on instantaneous changes in the weakly exogenous variables.

The existence of a single cointegrating relation where population, per-capita real income and real house prices are weakly exogenous also means that the long-run residual series constitutes the cycle component of residential investment in the trend-cycle decomposition proposed by Gonzalo and Granger (1995). The corresponding trend component, which is a linear combination of movements in population, per-capita income and real house prices, has been shown to be smoother than the trends which have been extracted by standard univariate filtering techniques. In particular, the boom-bust movement in residential construction after reunification is part to the Gonzalo-Granger cycle component while the Hodrick-Prescott and the Baxter-King filters assign it largely to the trend component.

A corollary of the approach to modelling residential investment within a vector error correction model is that the short-term dynamics of dwellings construction have been explained only insofar as they are governed by the adjustment process towards the long-run equilibrium. Together with government interference in the various segments of the housing market, this turns out to be a dominant feature in residential investment throughout the 1990s and the first half of the 2000s. The influence of other economic factors such as user cost of housing capital, lending conditions and the degree of affordability might be concealed. In any case, the integration of these elements in the specification of housing demand and, thus, residential investment was beyond the focus of this paper.

The outcomes of alternative trend-cycle decompositions can be assessed in terms of economic plausibility. However, the ultimate proof which variant comes closest to reality can never be offered as there are no direct observations of trend and cycle components. As a consequence, the critical statements on simple statistical filtering techniques are not meant to deny their practical relevance, in particular when the intention is to conduct a comparative analysis on a unique methodological basis (Álvarez et al., 2009, for instance). The example should rather remind users to be cautious in dealing with trend-adjusted figures in this context.

A Time series and unit root tests

The time series of real residential investment, population and national income are taken from quarterly national accounts, adjusted for seasonal and working-day variations. The quarterly data on (nominal) house prices are produced by interpolating the annual price series for new dwellings, specifically terraced houses and flats, which is constructed by the Deutsche Bundesbank on the basis of primary data collected by BulwienGesa. The deflators needed to calculate real national income and real house prices are computed using national accounts data.

The time series refer to Germany as a whole from 1991 onward. They are chained with the corresponding series for western Germany starting in 1975 and with an overlap in 1991. Real quantities and price indices refer to 1991 as the reference year. This choice ensures comparability between the west German and the pan-German subsamples. The time series are transformed in natural logarithms.

Table 4: Unit root tests

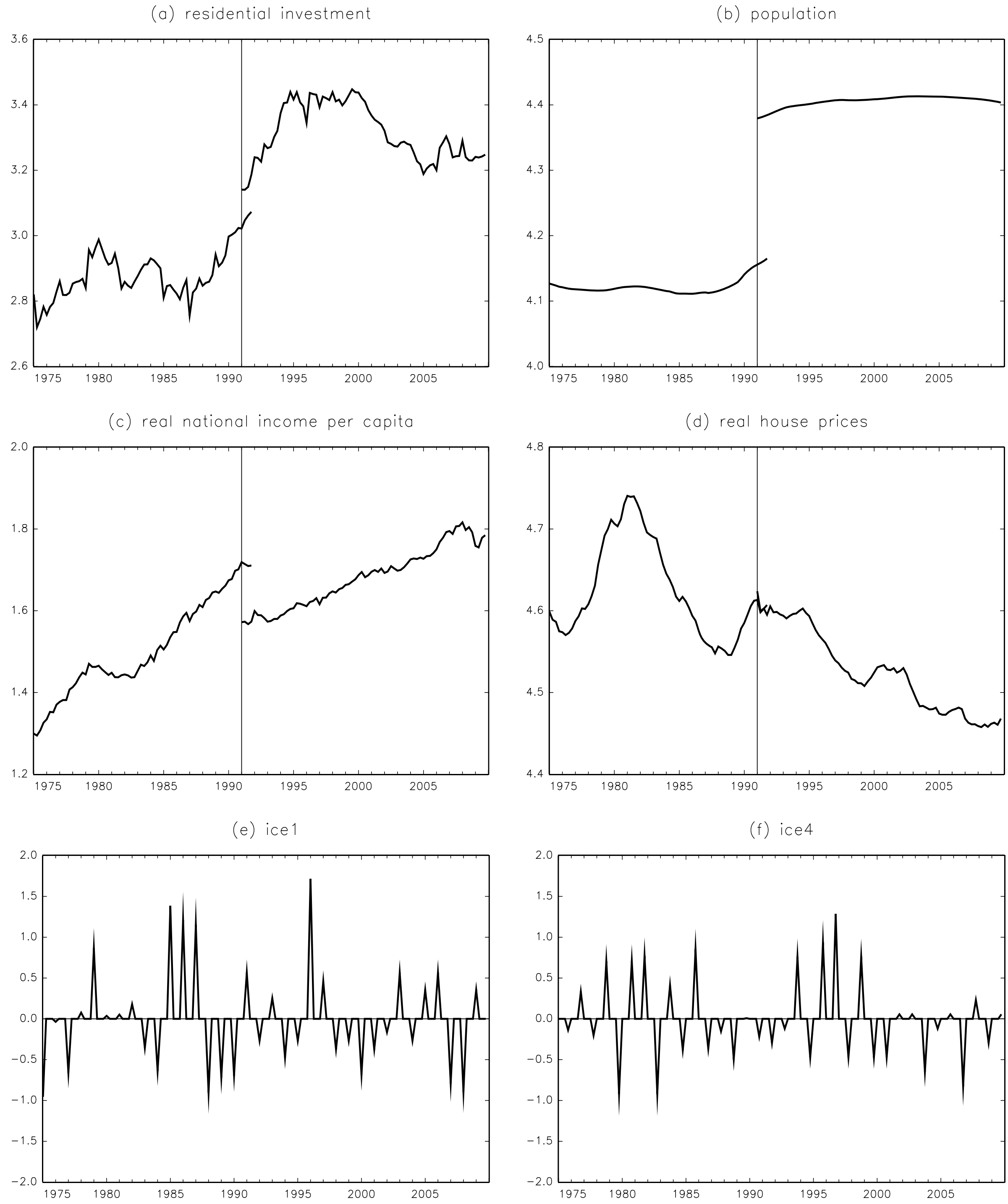
series	deterministic	ADF		PP		KPSS	
ri_t	$\mathsf{c, t, S(91{:}1)}$	(0)	−3.28	(3)	−3.29	(4)	0.211^{**}
pop_t	$\mathsf{c, t, S(91{:}1)}$	(0)	$-3.60^{(*)}$	(1)	$-3.69^{(*)}$	(4)	0.205^{**}
inc_t	$\mathsf{c, t, S(91{:}1)}$	(0)	−2.21	(5)	−2.36	(4)	$0.115^{(*)}$
phr_t	$\mathsf{c, t, S(91{:}1)}$	(1)	−2.92	(8)	−2.54	(4)	0.190^{*}

The numbers in parentheses indicate the lag length in the ADF procedure and the bandwidth parameter in the PP and KPSS procedures. The ADF and the PP tests are performed following Perron (1989), as a structural break has to be considered in the specification. Critical values are taken from Table IV.B of this paper which are −4.32, −3.76 and −3.46 in the given setup. The KPSS test is performed according to Kurozumi (2002); critical values are found in Table 1b: 0.204, 0.134 and 0.106.**,*,$^{(*)}$ mean rejection of the null hypothesis at the 1%, 5% and 10% level respectively.

Real residential investment, population, real national income per capita and real house prices are plotted in Figures 7(a) to (d). Visual inspection shows them to be nonstationary even if the drifts are only slightly positive in the case of residential investment and population. While per-capita income tends upward over the whole sample with a setback at reunification, real house prices have fluctuated substantially around an overall negative trend. The order of integration is checked by standard unit root tests, namely the augmented Dickey-Fuller (ADF) test, the Phillips-Perron (PP) test and the KPSS test proposed by Kwiatkowski et al. (1992). The ADF and the PP procedures test for a unit root under the null hypothesis. In the present setup, the null hypothesis of the KPSS test is trend-stationarity including a mean shift in the first quarter of 1991.

As Perron (1989) has suggested for the ADF and PP tests, the structural break is taken into account by a prior removal of the mean shift and by applying the critical values

Figure 7: Time series



The vertical line indicates the date from which data for Germany as a whole are available. As regards Figures (a) to (d), note that data for western Germany and Germany as a whole are available for a short overlapping period, namely the year 1991.

reported in this paper for $T_B/T = 0.5$ where T_B is the date of the break and T the length of the sample. Against the alternative of trend stationarity with a change in mean in the first quarter of 1991, the null hypothesis of a unit root cannot be rejected for any series at the 5% level. The evidence seems clear-cut except for population, where both the ADF and the PP procedures signal rejection of the unit root hypothesis at the 10% level. The mean shift in the KPSS test procedure is captured by the test structure proposed by Kurozumi (2002). For all series under consideration, the null hypothesis of stationarity is rejected at conventional significance levels, confirming the I(1) assumption for residential investment, population, real national income per capita and real house prices.

Figures 7(e) and (f) display the time series plots of the variables $\mathsf{ice1}_t$ and $\mathsf{ice4}_t$ which measure the effects of unusual weather conditions in the first and the fourth quarter. The time series are constructed on the basis of information about the number of days with a maximum temperature below freezing in the respective quarters. As the average impact of production impediments due to weather conditions in the first and fourth quarters is removed by seasonal adjustment, the variables are defined in relation to the long-run seasonal means; to be precise, as a percentage of the long-run average of days with a maximum temperature below freezing in the respective quarters.

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