Has the Euro Sustainably Increased Home Price Co-Movement?

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Abstract

A low level of co-movement between different euro country housing markets creates difficulties for the ECB in setting monetary policy. Such co-movement across euro zone countries has been the subject of a number of studies, using different methodologies and finding mixed results. In this study, we use endogenous break methods to explicitly test for whether the introduction of the euro has changed home value co-movement. We also employ informal correlation analysis. Endogenous break results indicate no sustainable increase in co-movement attributable to the euro, while correlation analysis is suggestive of a *decrease* in synchronization since the currency's introduction.

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JEL Classifications: R31, E32, E52, F4

1. INTRODUCTION

Housing co-movement across the euro zone is a topic of particular salience, as there is a single currency for all nations using the euro. Thus there is only one monetary policy set by the European Central Bank. If there is a high degree of co-movement in housing returns across member countries, the ECB can incorporate housing into its policy decisions. This could be important, as housing has been shown to be the most predictive sector for the macroeconomy, at least for the US (Leamer, 2007, 2015). Moreover, Claessens, Kose and Terrones (2012) find that recessions in the wake of housing downturns are "longer and deeper" than recessions not preceded by such episodes. Lastly, regardless of whether the ECB directly incorporates housing into its policy, a high degree of co-movement in housing makes a single monetary policy more "optimal" and suitable for the different nations of the euro. If nations' housing sectors are in different cyclical states, monetary policy-be it loose or tight-will be suitable for some countries but the exact opposite of what other countries would desire. For instance, if home prices are rising to bubble levels in say Spain, but suffering a downturn in France, a tight policy will be appropriate for Spain but possibly devastating for France. The opposite would hold true for loose policy.

Accordingly, a number of papers have been written on international house price co-movement in the euro zone. Alvarez, Bulligan, Cabrero, Ferrara and Stahl (2010), Ferrara and Koopman (2010), Van Steenkiste and Hiebert (2011) and Gupta, Andre and Gil-Arana (2015) all examine the interaction of house prices between euro zone countries.

Our goal in this paper is to determine whether the advent of the euro has sustainably *increased* co-movement among home prices in different euro countries. There is theory on the endogeneity of optimal currency area determinants, such as business cycle synchronization, as well as conjecture, that such a monetary union should increase house price synchronization. Previous papers have either examined whether countries which would eventually join the euro had high co-movement over periods which include years before and since the euro's introduction, or have tested for the impact of the common currency by splitting the sample at the beginning of the currency union.

In this paper, we first employ a recently developed data set of international house prices which overcomes some previous issues with cross-country house price comparisons by using a consistent methodology for all nations. We then examine correlations among eight different euro zone countries for years spanning 1975-2015. This period of course entails both pre- and post-euro years. To get a gauge of the euro's impact, we will split the sample at different points-first at 1989, which has been labeled as the first stage of the European Monetary Union, or EMU (Van Steenkiste and Hiebert, 2011), as well as before and after 1999, which is when the euro was launched, in order to see whether home price correlations rose or fell in response to monetary union.

Such correlation analysis is of course informal, as choosing break dates in 1989 or 1999 based on known policy change presents an endogenous break problem (Hansen, 1992). If one has knowledge of an economic or financial event, and obtains estimates before and after the event, one may find "significant" change where none has occurred.

We will thus seek to test for change with a method that allows for any such breaks to be discerned endogenously. We therefore first compute rolling unconditional correlations for country pairs as well as overall country average co-movement. We investigate both what might drive co-movement and whether the euro may have played a role in increasing these correlations. We then test for changes in these correlations with endogenous break tests, to see if breaks seem to coincide with euro adoption. There are of course tools beyond rolling unconditional correlations which are typically employed to study changes in asset price interaction. We thus estimate a DCC GARCH model for those six euro nations where significant ARCH effects are found. We then model the obtained, dynamic time-varying correlations as autoregressive processes. As with the unconditional correlations, we then test for change in these dynamic conditional correlations with endogenous break tests.

Measures using filtered data, not just changes, have been employed in previous studies of home price co-movement. We will accordingly utilize a measure that builds on and goes beyond previous measures of coherence and concordance that is known as Similarity. This metric takes account not just of which phase of the cycle different national home markets may be in but also differences in the amplitudes of the cycles. It is developed both for each country to examine its co-movement with all other national housing markets collectively and also for all countries to get an overall measure of how tightly connected are the housing markets of the euro zone.

To anticipate our results we find, contrary to some previously expressed expectations, that the correlation analysis yields no evidence that the euro has increased home price synchronization. Indeed, results for Germany, the largest economy, are suggestive of a divergence between home values among the euro members in response to currency union. For the rolling bilateral and national correlations, we find that proximity between nations has a positive-and homeownership rates and current account balances have a negative-impact on co-movement. A key finding regarding the common currency is that the break tests applied to the rolling correlations show no indication of an increase in co-movement due to the euro-in fact there is some indication that the euro, if anything, is associated with lower home price correlation.

Moreover, the DCC GARCH analysis yields mixed findings. There are some significant breaks in the dynamic conditional correlations in the 1990s which may indicate an increase in co-movement that on its face may seem attributable to the euro, while there are other breaks which suggest a decrease in co-movement at this time. What is clear is that any increase for a country's dynamic correlation which occurred around the time of the euro was not sustained, as the dynamic correlations fall subsequent to euro adoption.

For similarity, we find, analogous to the results using the DCC GARCH, that there appears to be an increase in co-movement around the time of the euro. However, this increase is not sustained, and overall similarity falls after euro adoption and ends at its lowest level of the sample. These results stand in contrast to previous findings.

This paper proceeds as follows. The next section details the previous literature. The third section describes the data and methodology. The fourth section presents our results, and the fifth concludes.

2. PREVIOUS LITERATURE

Growing globalization, especially in the form of cross-border financial investment but also trade flows may have an impact on home values. Ferrero (2015), Gete (2015), Miles (2019) and Punzi (2013) and all examine how home values interact with capital flows. This is of course of interest to investors or institutions with exposure to home values in different countries. In addition, housing has perhaps a greater impact on business cycles than any other industry (Leamer, 2007, 2015). Claessens, Kose and Terrones (2012) find recessions which are preceded by a housing downturn are more severe than contractions in which a housing crash had not previously occurred. This finding certainly has relevance for the experiences of Ireland, Spain and the US, which all had harsh recessions following home industry downturns.

Given the importance of housing to the macroeconomy, and the possible impact of global factors on home values, a research agenda on cross-border home price movements has developed. If co-movement between home values in different countries is substantial, knowledge of home values abroad could help in forecasting house prices at home. Moreover, if home prices are highly correlated across nations, this would make a common monetary policy more feasible. This would not have relevance to home price co-movements between, say the US and Japan, which are unlikely to form a monetary union anytime soon, but it has great importance to those European countries which share a common currency.

Would adoption of the euro be expected by some economists to increase synchronization among different national housing markets? Yes, because the euro was expected, by some economists, to increase the synchronization of business cycles across the monetary union (Frankel and Rose, 1996). Adopting the euro means adopting a common currency for all nations in the zone. When is it optimal to give up a national money to form a currency union? In his original "Theory of Optimal Currency Areas", Mundell (1961) pointed out that a currency union between different countries would be more likely beneficial if there was a high level of trade as well as labor mobility between them.

Others have pointed to a high level of business cycle synchronization as a key criterion for making a single money desirable. This can be very important. By joining a currency union, member countries given up their own central banks and ability to conduct independent monetary policy. If, say the Irish economy (and hence, likely the Irish housing market) is booming, while the German economy and housing market are stagnant, a single European Central Bank cannot provide proper policy for both nations. The loose monetary policy that would benefit Germany would fuel further bubbles and inflation in Ireland, while the tight policy that would help Ireland would be very harmful to Germany.

A high level of synchronization of the different national business cycles thus helps make a common currency more desirable. However, it has been argued that high levels of trade and business cycle synchronization, though seen as desirable before countries enter a monetary union, may be increased by the very act of entering the common currency. Rose (2000) presents results indicating that joining a currency union will increase-indeed triple-the level of trade. Higher trade can then increase business cycle synchronization. Of course greater trade may lead to greater specialization, which could drive business cycles apart rather than closer together. But Rose argues that if trade is predominantly intra-industry, it will raise business cycle synchronization. And Frankel and Rose (1998) present evidence for industrialized countries that greater trade is associated with more business cycle synchronization.

Thus Frankel and Rose argue that "the OCA (Optimal Currency Area) criteria are endogenous" (1998, p. 1010). Countries that may not seem suitable for a common currency prior to entering a monetary union could be made suitable just by joining the monetary union, in this line of reasoning. First, upon joining the euro, trade will strongly increase. Then this higher trade will make business cycles more synchronized.

As housing is tied to the business cycle (perhaps more than any other sector, i.e. Leamer, 2007, 2015) greater business cycle synchronization should increase housing co-movement. While not citing Frankel and Rose specifically, Van Steenkiste and Hiebert (2011) state "co-movement in house prices across countries may be particularly relevant in the euro area, given a general trend with monetary union toward increasing linkages in trade and financial markets" (p. 299). Similarly, Gupta, Andre and Gil-Alana (2015) when discussing house price co-movement, posit "synchronization could be even stronger in the euro area than at the OECD or global level, not least because member states share a common monetary policy that affects income developments and mortgage rates" (p. 3124).

The impact of a currency union on business cycle and housing co-movements is not as clear as stated by Frankel and Rose, however. First, other researchers have, upon using different methodologies, found no significant effect of a common currency on trade (see Persson, 2001). In addition, there are channels besides trade through which a monetary union can either increase or decrease synchronization. For instance, a fixed exchange rate regime (of which a common currency is the most rigid) can facilitate capital flows (Eichengreen and Hausmann, 1999, Obstfeld, Ostry and Qureshi, 2018). This could send capital from stagnant nations, say Germany in the early 2000s, into faster growing economies and housing markets, such as Ireland and Spain. These flows could then inflate bubbles in the recipient countries, driving business cycles apart. And when these bubbles burst, causing severe recessions in the capital receiving countries fluctuations can move still further apart. How the euro affects housing market co-movement is thus an empirical question.

The potential problems of having a common currency when business cycles are not well aligned has been a topic of research for the euro (Bayoumi and Eichengreen, 1993 is one of many examples). And again, the introduction of the common currency can exacerbate, rather than ameliorate this lack of alignment. DeHaan, Hessel and Gilbert (2015) discuss the persistent financial imbalances that developed between countries subsequent to the euro's introduction. Obstfeld (2013) describes the inadequacy of the eurozone's policy infrastructure for handling the aftermath of the financial crisis that began in Europe in 2009.

Given the potential impacts of the euro common currency on business cycles, trade and asset prices, several studies have been conducted on co-movement in housing values among member nations. Ferrara and Koopman (2010) note that "the existence of a common housing cycle among the countries of the zone could lead the ECB to integrate more easily the evolution of this specific asset price into its assessment. On the other hand, if country-specific cycles were too large, this would complicate the task of the ECB" (p. 4). The authors thus examine real house prices and GDP for four members of the zone-France, Germany, Italy and Spain. They utilize data spanning 1981-2008, and employ a multivariate unobserved components model. Results indicate that business cycles-GDP-appear strongly related for France, Italy and Spain. However, there is less evidence for synchronized housing cycles, although French and Spanish home values appear strongly related.

Alvarez, et al. (2010) examine the same four nations-France, Germany, Italy and Spain-with data from 1980 to 2008. These authors employ a number of measures, such as correlation and concordance indices and find, similar to Ferrara and Koopman that GDP is highly synchronized across countries but that this is not the case for housing. They then go on to investigate specifically whether the advent of the euro has led to an increase in co-movement for output and home values. They split the sample at 1998 and find that in the euro years most indicators suggest an increase in synchronization, including for nominal house prices. Real house prices, however, actually seem to have less co-movement since the common currency came into being.

Van Steenkiste and Hiebert (2011) gather data on house prices, real disposable income, and real interest rates over 1971-2007 for seven euro zone countries-Belgium, France, Germany, Ireland, Italy, The Netherlands and Portugal. The authors employ a global VAR and the impulse responses indicate that domestic home price developments do at times spill over into the housing markets of neighboring countries. The authors then split the sample at the end of 1989, a point which "closely corresponds to the onset of the first stage of the European Monetary Union" (p. 303). They only investigate whether interest rates have a greater impact on home values after 1989 than in previous years, not examining whether home price spillovers were greater or lesser before or after this break point. They do find that interest rates seem to have a larger effect on home prices in the latter years.

Gupta, Andre and Gil-Alana (2015) examine home price co-movement for the eight euro zone nations of Belgium, Finland, France, Germany, Ireland, Italy, The Netherlands and Spain. Their sample spans 1971-2012. They find cointegrating relationships among a number of their sample countries, but find German home values “seem to move in the opposite direction from other countries” (p. 3124), which the authors attribute to “capital flows associated with current account imbalances”. This is a very important point made by the authors. They cite papers on home prices and current account imbalances, and point out that “housing booms in Greece, Spain and Ireland were associated with a sharp deterioration in current account balances. At the same time, large external surpluses coincided with declining house prices in Germany” (p. 3126). These are important points, as a common currency could actually drive home prices further apart by sending capital from surplus countries like Germany to smaller deficit nations.

Also note that, in addition to facilitating capital flows, a common currency could facilitate labor migration across euro-zone countries. This migration could also have ambiguous effects on business cycle synchronization. Labor leaving a country in recession for a nation in expansion could increase co-movement. On the other hand, workers-hence taxpayers-fleeing a country in a debt crisis could exacerbate differences in fluctuations. For a discussion of migration issues in some euro-zone nations see Gonzalez and Ortega (2013), Schundeln (2014) and Granato, Haas, Haman and Niebuhr (2015).

These studies have all addressed an important topic. But only two test specifically for breaks of some kind among member housing markets which may indicate whether the euro itself has changed home price co-movement. And we want to conduct tests which allows for breaks to be determined endogenously, to avoid Hansen’s (1992) endogenous break problem. We thus set out the following methodologies.

3. DATA AND METHODOLOGY

In order to test whether the euro has induced an increase, or perhaps a decrease, in synchronization among member house price cycles, we start with a recently developed data set which has been calculated with a consistent methodology across nations. Previous studies on cross-country home price co-movement have been plagued with the issue of using different indices for different countries utilizing different methods, making comparison across borders questionable. Hirata, Kose, Otrok and Terrones (2013) far from claiming that their study is immune to such problems, acknowledge on page 8 that "House price series are subject to various problems given that different countries use different concepts to keep track of price movements in housing markets."

Mack and Garcia (2011) have developed a set of quarterly real, seasonally adjusted home price series using a consistent methodology that is consistent with the US Federal Housing Finance Agency (FHFA) index. This index has better coverage than the more publicized Case-Shiller index, which is frequently cited in the media. The FHFA index has better geographical coverage of US states and it, rather than the Case-Shiller, is employed by the US Federal Reserve when modelling the impact of home values. Thus the Mack and Garcia index is a big improvement over the previous method of collecting different indices based on different methodologies from different nations to examine home price co-movement. Our data set will initially consist of eight countries-Belgium, Finland, France, Germany, Italy, Ireland, The Netherlands and Spain. The data runs from the first quarter of 1975 to the first quarter of 2015 (Mack and Garcia update their data on the Federal Reserve Bank of Dallas website at https://www.dallasfed.org/institute/houseprice/).

Given the non-stationarity of asset prices, we will employ the log year-on-year difference for home values so that we have annual returns. We first as a preliminary exercise examine correlations between home value returns in the eight nations. We will then split the sample for two dates-1989 and 1999-the first corresponding to the end of the EMS period according to Von Hagen and Neumann (1994) as well as the date chosen by Van Steenkiste and Hiebert (2011), and second to the launch of the common currency, so to observe whether co-movement among countries increases or decreases in response to monetary union.

Splitting a sample and examining whether there has been a "change" in some estimator can lead to falsely concluding that there has been a significant break of some sort though in fact none has occurred. Hansen (1992) explains that since the date of alleged change is chosen because the researcher, who knows the data, has suspicions that a certain observation was a point of parameter change, one is likely to reject a null hypothesis of no change, using standard critical values, if one is conducting a formal test. We thus interpret the results of the correlation exercise as very tentative.

We next seek to examine more dynamic measures of co-movement. We will thus calculate twelve quarter rolling bilateral correlations for all country pairs. We will also average these bilateral correlations for each country to get a rolling unconditional measure of each national housing market vis-à-vis the other seven countries collectively.

We will then examine whether national housing market co-movement seems to be related to certain economic attributes, such as country proximity, trade between countries, financial development, financial systems (bank-based versus market-based), homeownership rates, rental market assistance policies and the current account balance. To anticipate our results, we find that proximity is positively related to co-movement while homeownership rates and the current account balance appear negatively related with house price correlation.

Given the number of countries (eight), this type of analysis cannot be formal. We will thus formally test for changes in the dynamic, unconditional measures of co-movement, and see if these changes are related to the euro. To do so, we will model the dynamic correlations as autoregressive processes, with the number of AR lags chosen by the SIC criterion. We will then test for breaks with the Lee-Strazicich method. The Lee-Strazicich procedure both tests for a unit root and structural breaks in a series. This is important, as the presence of breaks which are not empirically modelled lowers the power of standard unit root tests. And of course we seek to find any structural breaks, to see if structural change is related to the euro. The Lee-Strazicich technique, unlike some other unit root tests, allows for breaks both under the null hypothesis of nonstationarity and the alternative hypothesis of stationarity. This is a major improvement over previous methods, which only allowed for breaks under the alternative hypothesis of stationarity.

We do seek to formally test for change using measures besides unconditional correlation which have been used in modelling asset prices that also allow for endogenous breaks. We thus employ the Dynamic Conditional Correlation (DCC) GARCH model of Engle (2002). This method models each return series as an AR GARCH process, and then calculates bilateral correlations between all series in question for each quarter. A new correlation is calculated for each quarter, hence they vary through time, and we get a series of dynamic conditional correlations (DCCs). These DCCs can be modeled as AR processes.

The DCCs may be (and to anticipate our results, often will be) non-stationary. We thus test for structural change with a unit root test that allows for endogenous breaks. As with the unconditional correlations, we will employ the Lee and Strazicich method. We note that the GARCH methodology has been applied to financial assets such as equities (Glosten, Jagannathan and Runkle, 1993, is one of countless examples). Houses are also an asset, and GARCH models have been employed in investigating housing returns (Miles, 2008 is an early example). Indeed, the specific DCC-GARCH estimator we will use has been utilized to examine housing market dynamics across US metropolitan regions (Zimmer, 2015). Thus GARCH is a tool which has frequently been applied in studies of housing.

At the same time, housing markets do not behave exactly like equity markets. Housing is a consumption good in addition to being an asset, and construction of new houses takes time. Unsurprisingly, house prices don’t change as quickly as equity values. Case and Shiller’s (1989) classic study showed that house prices did not follow a random walk, as equity prices have been shown to do

(See also Hessel and Peeters (2011) for a discussion of the persistence and slow adjustment of house prices).

Thus while GARCH has been applied to housing markets, we will also employ an alternative method which has been applied less frequently to equities but more often to housing (Miles, 2017b).

This measure, known as similarity, was developed by Mink, Jacobs and DeHaan (2012). These authors were examining the coherence of output gaps in the euro zone. The authors point out that previous measures of coherence, such as concordance indices (Harding and Pagan, 2002), which measure how often a group of variables are in the same phase of the cycle, and those correlation-based metrics employed by Flood and Rose (2010) fail to account for the fact that the output (or in our case, housing prices) can be highly "correlated" but still exhibit large cyclical differences. Mink, et al. point out that there could be large differences in amplitude between the cycles of countries, even if they are highly (even perfectly) correlated. For instance, if housing cycles are very volatile in say, Spain, but milder in Germany, the best monetary policy if both Germany and Spain were in an expansion would be much tighter for Spain than it would for Germany (although tight money would be appropriate for both countries).

Mink, et al. fortunately have come up with a metric that takes account not just of whether two countries are in different phases of the business cycle but also how different the amplitudes are. To obtain the cycle of a series, the authors filter the data with the Christiano-Fitzgerald technique. The Christiano-Fitzgerald method is the best filter for such purposes, relative to earlier methods like the Hodrick-Prescott and Baxter-King filters. The filtered, cyclical portion of a country's output (or house price) at period t is denoted as $g_i(t)$. To calculate similarity, a reference cycle is required. Mink, et al. use the median of all country cycles as the reference. The reference cycle at period t is $g_r(t)$. The similarity measure for a given country at time t is:

$$\gamma_{it}(t) = 1 - (| g_i(t) - g_r(t)|) / \sum_{i=1}^{n} | g_i(t)|/\text{n}, \qquad (1)$$

For a given country, similarity in a period can range from 1-n to a maximum of positive one. When similarity equals 1, the cycle of the given country is exactly the same as that of the reference. There is an aggregate measure of similarity for all sample countries at time t:

$$\gamma(t) = 1 - (\sum_{i=1}^{n} | g_i(t) - g_r(t)|) / \sum_{i=1}^{n} | g_i(t)| , (2)$$

This metric ranges from zero to one. If, in a given period, overall similarity is one, then all countries in the sample are having the same identical cycle.

As with the rolling conditional and unconditional correlations, we will model the country-specific and overall similarity measures as AR processes, and then test for endogenous breaks that may or may not correspond to the euro with the Lee-Strazicich method.

4. RESULTS

We first as a preliminary exercise examine correlations between home value returns in eight of the euro zone nations for which we have solid, consistent data from Mack and Garcia. These happen to be the same eight nations employed in Gupta, et al. (2015). Levels of home prices have been shown to be non-stationary (Gupta, et al. 2015 among others) so we employ the log year-on-year difference for returns. The sample for returns thus runs from 1976:1 to 2015:1. The summary statistics are shown in Table 1. Although the returns appear to cross their own means reasonably often, to be absolutely sure the return data are stationary, we apply the ADF unit root test. Table 2 displays the results, which indicate we can reject the null of a unit root at the five percent level in all cases.

Table 3 displays results for the correlations over the entire 1976:1-2015:1 sample. As there are eight nations, there are twenty-eight different correlation coefficients. Nineteen of the correlations for the whole sample are positive, while nine are negative. For comparison, while Gupta, et al. employed different sources for their house price indices and used a slightly different span of time (1971-2012) they found six cases of negative correlation of house price returns (see Gupta, et al., p. 3130). Our most negative correlation comes between Germany and Finland, at -0.397. Interestingly, Germany and Finland also had the most negative correlation in Gupta, et al's sample, although the magnitude was not as high (-0.283) in their results. In our sample, the highest, most positive correlation was between Belgium and neighboring Netherlands at 0.694.

The nine country pairs with negative house return correlations are Belgium/Finland, Belgium/Italy, Finland/Germany, Finland/Netherlands, France/Germany, Germany/Ireland, Germany /Italy, Germany /Spain, and Italy/Netherlands. Germany thus has five negative correlations, more than any other country. This qualitatively matches the findings of Gupta, et al. who found that in their analysis, "Prices in Germany seem to move in the opposite direction from other countries, which may be related to capital flows associated with current account imbalances" (p. 3123).

To get a very rough sense of how monetary union may have affected co-movement we next split the sample into two, with the first half consisting of the quarters from 1976:1 through 1989, and the second half consisting of the remainder quarters from 1990:1 through 2015:1. As explained this amounts to imposing an exogenous break at a known point and so results are at best only very tentative. The "breakpoint" at the end of 1989 corresponds to that chosen by Van Steenkiste and Hiebert (2011). Results for the 1976:1-1989:4 sample are displayed in Table 4. Interestingly, only seven of twenty-eight possible pairs are negative before 1990, as opposed to nine such instances in the whole sample. The lowest, i.e. most negative, correlation was between The Netherlands and Italy at -0.489, followed closely by Germany/Finland. The highest correlation was once again between Belgium and The Netherlands. Finland is the only country with which Germany displays negative co-movement in these years. The other six negative correlations are Belgium/Finland, Belgium/Italy, Finland/France, Finland/Ireland, Finland/Netherlands, and Netherlands/Italy. Thus Germany's negative relationship with France, Ireland, Italy and Spain for the entire sample is absent when *excluding* the years close to the euro. This result is confirmed in Table 5, which shows correlations when the sample is restricted to 1990:1-2015:1. There were again seven negative correlation pairs, but all were with Germany.

Given that the euro did not fully commence until 1999, we follow Alvarez, et al. (2010) and split the sample for the correlations into the 1976:1-1998:4 quarters and 1999:1-2015:1. Results for the former are displayed in Table 6. The results are somewhat similar to those found using 1990 as the breakpoint. In the pre-euro sample, only five of twenty-eight possible correlations are negative. The lowest is Netherlands/Italy, with a value of -0.476. The other negative pairs are Belgium/Finland, Belgium/Italy, Finland/Germany, and Finland/Netherlands. Again, *before* the euro, Germany had negative co-movement in housing returns with only one country.

In contrast, in the euro years, Table 7 indicates that Germany again exhibits more negative co-movement than when countries retained their own currencies. There are seven negative correlations, and all involve Germany. Indeed, Germany has a correlation of -0.632 with neighboring France and of -0.621 with neighboring Belgium. Thus it again appears that the advent of the euro led to less, rather than greater co-movement between the largest economy Germany and other countries in the currency union.

We note that in the pre-euro years, Finland had a very low level of co-movement vis-à-vis other housing markets. Finland's circumstances make it something of an outlier for much of this period. Miyagawa and Morita (2009) provide a detailed analysis of the situation. These authors point out that, starting in the mid-1980s, Finland experienced an economic and financial boom. First, the price of oil, an import fell, and the price of forestry products-a major export-rose. This change in the terms of trade was accompanied by financial liberalization, a large capital inflow and a deteriorating current account. Lending for property rose, as did asset prices.

The bursting of the bubble in 1990 was prompted in part by the collapse of the Soviet Union, a key trading partner. The recession that followed in the 1990s was harsh, as is typically the case for recessions that follow credit booms and asset price bubbles. In this context it is thus not surprising that Finland's housing displayed little co-movement with more moderately growing national home markets in these years.

In order to examine the dynamics of co-movement, we next turn to the twelve-quarter rolling unconditional bilateral correlations. Looking at Table 8, which ranks these bilateral correlations over the whole sample, one thing that stands out is that to some extent proximity is associated with co-movement. The top three of the 28 pairs-Belgium/Netherlands, France/Spain, and Belgium/France-are all contiguous, while the lowest two, Finland/Germany and Italy/Netherlands-are fairly far apart. It is not clear, however, exactly why proximity would necessarily lead to greater co-movement. One possibility is trade ties. The gravity model would suggest that proximity would lead to greater trade. Of course, as previously discussed, greater trade would have an ambiguous effect on home price co-movement.

It is true that for Belgium and the Netherlands, which exhibit the greatest correlation, there is a fairly high level of bilateral trade (data on trade is available at the World Bank (2018) website World Integrated Trade Solution). The level is not overwhelming, however. In the case of the Netherlands, Belgium is the second leading trade partner in terms of both imports and exports. However, Germany is the Netherland's top trading partner in terms of both imports and exports, and as Table 8 shows, Germany and the Netherlands have a fairly low level of bilateral correlation, ranking 19th of 28 possible cases. For Belgium, the Netherlands is the top source of imports and third largest export destination. Germany is the top source of Belgium's exports and the second largest source of imports, but Belgium and Germany have palpably less co-movement (ranking 13th of 28 possible cases) than Belgium and the Netherlands.

For the second highest bilateral correlation, between France and Spain, it is true that France is the top export destination for Spain, and second largest source of imports. But while Spain is the second largest source of exports for France, it is not even in the top three sources of imports. In the case of the third highest correlation (Belgium and France), Belgium is not one of the top three sources of imports or exports for France. France is only the second highest export destination for Belgian exports, and the third highest source of imports.

The impact of trade on co-movement is further cast into doubt when one examines the two countries with the lowest bilateral co-movement, Finland and Germany. Despite (or perhaps because, given the possibly de-synchronizing effects of trade) the negative overall correlation between the two nation's housing markets, Germany is the top trading partner for Finland-the biggest source of both imports and exports. So while proximity does seem to have some positive association with co-movement of housing, trade does not seem to be a positive determinant of correlation.

Given the possible disruptions of the global financial crisis, we also compute bilateral correlations for just the 2008:4-2015:1 period. Results are displayed in Table 9. There are some differences between co-movement over the whole sample and the crisis years. The big changes include the Italy/Netherlands correlation moving from next-to-last, and negative, to the highest in the post-crisis years. The Netherlands and Spain move from fifteenth highest to the second when the sample is restricted to the post-2008 quarters. However, results still reflect some key patterns-Germany exhibits generally negative co-movement with its euro zone neighbors, and Finland also has low co-movement with other countries, while Belgium still shows mostly greater co-movement with other nations' housing markets than is obtained by other countries.

Going beyond the bilateral co-movement, Table 10 shows the average of the bilateral correlations for each country. This is calculated by averaging all seven of each nation's bilateral correlations. Unsurprisingly, in light of the results of Tables three through nine, Germany displays the lowest overall co-movement with other euro-zone housing markets, while Finland and The Netherlands also have low correlations with other markets. The other countries' average correlations, from highest to lowest, are those of France, Spain, Belgium, Ireland and Italy.

Mirroring the results from splitting the sample in Tables 4 through 7, the second set of columns of Table 10, which restricts the sample to 1976:1-1989:1, shows that Germany has much higher correlation with other nations before important movements toward the euro. Results for the other countries are not especially different although Italy displays lower co-movement in these early years than in the full sample. For the post-1989 years, the third set of columns of Table 10 shows results that are nearly identical to those of the full sample.

Restricting the sample to the pre-euro years of 1976:1-1998:4 in the first columns of Table 10 (continued) results in similar estimates to those of the pre-1990 period. Germany is less an outlier than in the euro period, and Italy's returns exhibit less association with other countries than they would later. The next column, for the euro regime shows Germany's housing market truly seems to move in the opposite direction of its neighbors. And as with the bilateral correlations, we estimate average correlations for the global financial crisis beginning in the fourth quarter of 2008. As displayed in the last columns of Table 10 (continued), results are largely similar to those of the overall sample, although Italy again shows less co-movement than in other periods.

We next investigate whether the results suggest a pattern in co-movement as being related to economic differences between countries. Table 11 shows a ranking of bank-based versus market-based financial systems for the eight countries, based on Levine's (2002) classification. No clear pattern emerges. Finland, which ranks last in terms of market-based systems (and hence first in terms of a bank-based system) has the lowest overall correlation, which might suggest that more market-based systems lead to greater co-movement in housing. However, Ireland, which has the most market-based of all financial systems, has only a mid-level overall house price correlation. And France and Spain rank high in house price correlation but only in the middle of the market-versus-bank rankings. Similarly, the third column of Table 11 shows rankings for overall financial development, also taken from Levine (2002). Again, no clear pattern emerges. The Netherlands and Germany are the top-ranked countries in terms of financial development, which might suggest an inverse relationship between home value correlation and financial depth. However, Belgium is ranked lowest in terms of financial development but is third highest in overall correlation. Ireland is third highest in financial development and ranks fourth in overall correlation.

In the fourth column of Table 11, we show homeownership rates for the eight countries (Source: Trading Economics (2018) website https://tradingeconomics.com/country-list/home-ownership-rate). Although not perfect, there does appear to be a rough negative relationship between housing market co-movement and home ownership. Germany has by far the lowest homeownership rate and lowest overall home price correlation, while Spain and Belgium have high correlations and high homeownership rates. Again, the relationship is not perfect; France has a low homeownership rate but a high level of overall correlation. However, based on this sample, it seems that high homeownership rates are also associated with high co-movement.

While the fourth column of Table 11 shows the percentage of residents that are homeowners in each country-the remaining percentage of residents count as renters-the fifth column displays which percentage of renters in each country receive assistance in the form of free or subsidized rent. The Netherlands tops the list, with three quarters of residential tenants obtaining some form of assistance. Germany, which has the lowest homeownership rate, has the lowest proportion of tenants getting government help. While Germany has the lowest portion of renters getting help, and the lowest unconditional correlation, there does not appear, overall, to be a palpable relationship between rental assistance and co-movement. Finland, for instance, ranks as third highest in the portion of renters getting free or reduced housing, but next-to-last in co-movement. Belgium ranks third in co-movement, but next-to-last in the portion of renters getting assistance. Overall, besides Germany, the percentage of renters getting assistance doesn’t appear related to co-movement.

A clearer relationship is that between the current account and home value co-movement. The last column of Table 11 shows average current account balances. The top three countries in terms of current account balance (surplus)-The Netherlands, Finland and Germany-are the lowest in terms of house price correlation. Spain, the country with the second-highest overall correlation, has the highest current account deficit. These results bolster the case posited by Gupta, et al. (2015) that Germany's current account surplus helps keep its house price co-movement low with that of other countries in the euro zone.

Overall, it does appear that home price co-movement is positively related to country proximity, as well as negatively related to the current account balance and the homeownership rate. Correlations do not appear to be connected to levels of bilateral trade, financial development, whether a country's financial system is bank or market based, or the level of assistance to renters.

The analysis in Table 11 is necessarily informal due to the sample size. The relative lack of a relationship between house price correlation and financial systems or renter assistance does not necessarily imply these variables are unimportant for housing markets, or the differences or similarities among them in different countries. But these factors may play more of a role in explaining differences in the levels of home values across countries rather than correlation.

In twelve of twenty-eight cases, the values of the time-varying correlations at the end of the sample are actually negative. In some of Germany's bilateral correlations there is a sharp upward movement but only in the last couple of years, well after the euro was introduced. But to formally test for changes we will model the rolling correlations as autoregressive processes and test for endogenous breaks using the Lee-Strazicich method. As the correlations do not appear to be trending, we will use the "Crash" specification, with no linear trend, and choose the number of autoregressive lags with the SIC criteria. Results are shown in Table 12.

It is perhaps unsurprising, given that the correlations are bounded between minus one and one, that in most cases, the null hypothesis of a unit root is rejected. The ten percent, five percent and one percent critical values for the unit root test are -3.504, -3.842 and -4.545, respectively. Thus for only four country pairs-Belgium/France, Finland/Ireland, Finland/Spain and Germany/Spain-is the null of nonstationarity not rejected. In two other cases-Belgium/Germany and Finland/Italy-the null is rejected at the ten percent level, and in all other twenty-two cases the null is rejected at the five percent level.

There were fifteen significant breaks. None indicate a role for the euro in increasing house price co-movement. There is a break at 1998:1 for the Belgium/France correlation, which was only a year before the euro. However, note that breaks can be either positive or *negative*; that is, they could indicate a decrease as well as an increase, respectively, in correlation. And the 1998:1 break is negative, suggesting a decrease in co-movement just as the euro was close to being launched. There are also two breaks at 1990:2, both involving the Netherlands and its correlations with Ireland and Italy. As 1989 was identified as the start of the EMU by Van Steenkiste and Hiebert, these breaks in 1990 could be taken as suggestive of a synchronizing effect by the movement toward a common currency. But these breaks are also negative.

There are four breaks ranging from 1985 to 1987, two of which are positive while the other two are negative. There are four breaks from 2005 to 2008, three of which are positive and three of which are negative. Overall, then there is no evidence from these rolling correlations that the euro has increased home value co-movement. Indeed, the breaks from 1990 and 1998 would, if attributable to the euro, indicate, if anything, a reduction in correlation from a single currency.

We next aggregate the rolling measures to observe how they vary through time, and test for any changes related to the euro. The aggregate rolling correlations-computed as the average, for each country, of all seven bilateral correlations, are displayed in Figures 1 through 8. As with the bilateral rolling correlations, sustained upward trends are not readily apparent. The rolling correlation for France does appear to rise from the late 1970s to the late 1980s, but then it fluctuates for the rest of the sample.

We next apply the Lee-Strazicich test to the average rolling correlations, with results displayed in Table 13. Once again, the “Crash” specification is employed, given a lack of sustained trends. Results for the unit root tests indicate the null of nonstationarity can be rejected at the five percent level for all countries except Finland, where the null can be rejected at the ten percent level. There are nine significant breaks. Four of the breaks are between 1985 and 1987, and all are positive. There is one break at 2005:2, for the Netherlands, and it is negative. The only breaks which might be clearly related to the euro are breaks at 1990:2 for France, one at 1991:4 for The Netherlands, and one at 1992:2 for Germany. But all of these three latter breaks are negative. The aggregate rolling unconditional correlations are not suggestive of any increase in co-movement attributable to the euro.

Given the use of the ARCH model in housing studies, we next seek results for conditional correlations. Toward that end of obtaining DCC GARCH results, we first estimate autoregressive models for each of the eight country housing returns based on the SIC criteria, and test for ARCH effects. Table 14 displays the results. The null of no ARCH effects can be rejected in six of eight cases. It cannot be rejected for Germany and Ireland. Thus we will estimate DCC GARCH models for Belgium, Finland, France, Italy, The Netherlands and Spain.

For the conditional mean models we fit AR specifications, with the number of lags for each country determined by the SIC criteria. We then estimated a standard GARCH(1,1) specification for all six nations, using the BFGS method. Results are displayed for the conditional mean and conditional variance in Tables 15 and 16.

In the conditional variance, Belgium has an estimated "a" coefficient which is significant at the ten, indeed almost the five, percent level. Finland has a β estimate which is significant at five percent. For the other four countries all of the a and β coefficients are significant at less than the five percent level. We then follow the method described by Engle (2002) and obtain dynamic correlations between all six nations. Note that they are calculated from 1977:1 through 2015:1 as there were four lags in The Netherlands conditional mean model.

As with the unconditional correlations, we model the DCCs as autoregressive processes and test for breaks with the Lee-Strazicich method. The number of lags is again chosen by the SIC criteria, and given the possibility of trends in the DCCs, we added a linear trend if it was significant in the AR specification. The LS procedure allows for two kinds of breaks-a change in the intercept and a change in the trend. If the model contains no linear trend, the breaks are strictly in the intercept and such a specification is referred to as the "crash" model, which we have employed up to now. If there is a linear trend, this is known as the "break" model. This procedure allows for breaks in both the intercept and the linear trend.

Results for the break tests are shown in Table 17. It is perhaps not surprising, given that each DCC is bounded between -1 and +1, that in only three of fifteen possible cases (Belgium/Spain, Italy/Spain, and Netherlands/Spain) are there significant linear trends. We therefore tested these using the "break" specification (Model C in Lee and Strazicich). In the other twelve cases the "crash" model will be used.

Despite being bounded between -1 and +1, concerns over non-stationarity are well-placed. In only four of fifteen cases is there even a question that the DCC may be stationary; in all other cases there are clearly unit roots as the null hypothesis cannot be rejected at any standard level (the ten percent, five percent and one percent critical values for the crash model are -3.504, -3.842 and -4.545, respectively, and the corresponding critical values for the break specification are -4.989, -5.286 and -5.823). Only for the DCC between Finland and Italy can we reject the null of a unit root, and only at the ten, rather than the five percent level. In three other cases-Finland and the Netherlands, France and the Netherlands and the Netherlands and Spain-we can almost, but not quite reject the null of a unit root at the ten percent level. Again, in no case can we reject the null at the five percent level, despite employing a test which allows for breaks and thus helps overcome the low power of the ADF method.

As discussed, the breaks for the *unconditional* rolling bilateral correlations did not yield much evidence in favor of a synchronizing effect from the euro. Of the fifteen significant breaks, only three were in the 1990s, and all were negative.

In contrast, the structural breaks obtained from the GARCH model-the dynamic conditional correlations-do seem, at first glance, to be suggestive of a positive impact on home price synchronization. There are twenty-two significant breaks overall for the DCCs. Four are in the 1980s, years designated by Von Neumann and Hagen as the EMS period, and all are positive. There are also three breaks in the late 2000s.

More significantly, there are fifteen breaks in the 1990s. Six of these breaks are in the early 1990s, running from 1992:1 to 1993:2-the years of the ERM crisis and adjustment in response to the crisis. Three of these six breaks are positive. For instance, the Belgium/France DCC had a positive break at 1992:1, as did the DCCs for Finland/France and Finland/Netherlands. There were the nine breaks that occurred in the latter half of the decade, running from 1996:1 to 1999:4, spanning the run-up to and first year of the common currency. All but one of these breaks is positive. The Belgium/Finland, Belgium/France and Belgium/Italy DCCs had breaks at 1997:2 in the first two cases and 1998:1 in the latter. The dynamic correlation for France and Spain also exhibits a positive structural change at 1997:2. Thus at first glance it seems the euro may have had a role in improving co-movement.

However, even if one attributes all of these breaks to the euro (and that would likely be unwarranted), these breaks did not presage a lasting increase in dynamic correlation. For instance, all of the DCCs involving Belgium with the exception of Spain, as well as those for Finland/France and Finland/Netherlands would end the sample with values below those at the launch of the euro. Indeed, in eight of fifteen cases, the ending values were lower than those observed at the beginning of the common currency. Moreover, in eight cases the DCCs ended lower than their values at the end of 1989, the date chosen by Van Steenkiste and Hiebert as the start of European Monetary Union. In addition, all of the DCCs end below their *pre-euro* peak values. Thus even if the euro had some positive impact on co-movement, the effect has been transitory.

Finally, given its use in other studies of co-movement, we turn to results from the Similarity metric, which was calculated for each of the eight countries with data filtered by the Christiano-Fitzgerald method according to equation 1. The overall similarity measure was also calculated with filtered data, as per equation 2. Graphs are displayed in Figures 9 to 17.

We again model the metric as an autoregressive process, with the number of lags chosen by the SIC criterion, and a linear trend added if it was significant in the AR equation. Next, the Lee-Strazicich unit root test is performed to determine whether a series is stationary and to test for structural change. The test statistics displayed in Table 18 indicate we can reject the null hypothesis of nonstationarity in all cases.

Most of the significant structural breaks are not indicative of a synchronizing impact of the euro. Spain has one break at 1987:3, which is positive, while Finland has a negative break at 1986:1. For the years fairly close to the launch of the euro, France has two breaks at 1996:1, Germany has one break at 1994:1, Italy at 1993:4, and The Netherlands at 1995:2. In all cases, save for one of the breaks for France, these structural changes are negative. In addition, examining the national similarity graphs also provides little evidence that the common currency drove housing prices together. Belgium exhibits an increase in the years just before the euro, but its similarity subsequently falls and fluctuates, ending below its pre-euro highs. The same is true for Finland. Germany's similarity fluctuates with no discernible upward trend. Ireland's ends well below its pre-euro, and overall, average.

Italy's similarity has a slight upward trend since the mid-1990s, but also includes a pronounced decline in the mid-2000s and ends well below its pre-euro peak. The Netherlands shows no sustained increase in similarity, and Spain's does increase in the 1980s but fluctuates since. Overall, these national similarity metrics provide no evidence in favor of an increase in co-movement attributable to the euro.

However, total similarity has a significant break at 1999:4, shortly after the introduction of the euro. And examining Figure 9, there is a sample peak at 2000:1, just subsequent to the break. This again could be taken as evidence that the euro has had a positive effect on house price co-movement in the euro zone. However, even if this break is attributable to the euro, it is clear that the impact has not been sustained. Figure 9 shows that there has been a secular *decline* in total similarity since the peak just after the euro's introduction and total similarity ends at near zero, its lowest possible value.

5. CONCLUSION

The euro has been a source of controversy, as some have claimed it would increase economic integration by boosting trade, and then increase business cycle co-movement across member countries. It may also in principle have had an impact on co-movement of asset prices, such as home values.

Upon examining the issue of housing co-movement with the new Mack and Garcia dataset, we find no sustainable increase in co-movement associated with the common currency. Indeed, the informal analysis with correlations, while not definitive, is suggestive of a decrease in such co-movement, and results from the DCC GARCH and similarity indicate that any increase in co-movement which may have seemed attributable to the euro has not been sustained. These results contradict some claims made by authors of previous studies on housing in the euro. However, an examination of the potential impact of a single money on asset co-movement indicates the effect could be either to increase or decrease co-movement, just as it may increase or decrease business cycle synchronization. This does not make monetary policy in the euro zone easier. On the other hand, an upside to these findings for investors is that the euro, by not increasing home price co-movement, has not led to fewer portfolio diversification opportunities for individuals and institutions exposed to home values across different countries in the currency union.

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	Mean	SD	Max	Date	Min	Date	Skewness	P-value	Kurt.	P-Value	JB	P-value

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Table 1

Summary Statistics for National Annual Returns

Bel	2.42	5.6	13.8	1976:3	-15	1981:1	-0.91	0.000	1.21	0.002	31.5	0.000
Fin	1.67	7.8	21.4	1989:1	-20	1992:2	-0.59	0.00	0.89	0.02	14.5	0.000
Fran	2.23	5.06	13	2005:2	-7.8	2009:2	0.12	0.52	-0.8	0.02	5.56	0.06
Germ	-0.05	2.4	6.6	1978:1	-5.6	1984:2	0.399	0.042	0.09	0.82	4.23	0.12
Ire	2.84	8.89	23.3	1998:3	-19	2011:4	-0.11	0.54	-0.17	0.66	0.56	0.753
Italy	0.597	2.23	7.83	2004:3	-7.4	2008:3	-0.37	0.055	2.02	0.000	30.4	0.000
Neth	1.81	9.02	30.5	1977:2	-24	1982:1	0.108	0.58	1.37	0.000	12.63	0.001
Spain	-0.79	11.4	19.6	1988:1	-47	1982:4	-1.06	0.000	1.97	0.000	55.19	0.000

Mean and SD refer to the average and standard deviation of each region's annual return over the sample. Max refers to the maximum return for each respective region over the whole sample; the date to the right of Max is the date on which this maximum occurred. Similarly, Min and the date to the right of Min refer to the minimum return and the date on which this minimum return occurred. Skewness, Kurt. and JB refer to tests for skewness, excess kurtosis and normality (JB is an abbreviation for the Jarque-Bera test) and the P-value columns are the probability values for these respective tests.

Table 2
Unit Root Tests
Housing Returns

ADF	Test Stat.

BEL	-3.193*
FIN	-3.755**
FRA	-3.498**
GERM	-2.963*
IRE	-3.221*
ITA	-4.083**
NETH	-4.322**
SPA	-4.289**

One asterisk denotes rejection of the null at the five percent level, while two asterisks indicate rejection at the one percent level.

Table 3

Correlations between Housing Returns

	Belgium	Finland	France	Germany	Ireland	Italy	Netherlands	Spain
Belgium	1.000000	-	-	-	-	-	-	-
Finland	-0.12319	1.000000	-	-	-	-	-	-
France	0.571868	0.383533	1.000000	-	-	-	-	-
Germany	0.246334	-0.3972	-0.08766	1.000000	-	-	-	-
Ireland	0.393475	0.229770	0.400188	-0.015778	1.000000	-	-	-
Italy	-0.06338	0.241462	0.418633	-0.169365	0.217898	1.000000	-	-
Neth.	0.69425	-0.0913	0.33082	0.127216	0.4152	-0.26473	1.000000	-
Spain	0.38878	0.1842	0.57904	-0.139700	0.43847	0.45369	0.236705	1.000000

This table shows the correlations between country housing returns for the entire 1976:1-2015:1 sample

Table 4

Correlations between Housing Returns 1976:1-1989:4

	Belgium	Finland	France	Germany	Ireland	Italy	Netherlands	Spain

Belgium	1.000000	-	-	-	-	-	-	-
Finland	-0.33793	1.000000	-	-	-	-	-	-
France	0.740518	-0.046680	1.000000	-	-	-	-	-
Germany	0.551711	-0.469705	0.433764	1.000000	-	-	-	-
Ireland	0.599200	-0.229737	0.391750	0.745435	1.000000	-	-	-
Italy	-0.30703	0.499311	0.155925	0.053000	0.056470	1.000000	-	-
Neth.	0.794681	-0.336140	0.513140	0.370958	0.272907	-0.48937	1.000000	-
Spain	0.305952	0.082762	0.495228	0.109898	0.343898	0.314282	0.156784	1.000000

This table shows the correlations between country housing returns for the 1976:1-1989:4 sample

Table 5

Correlations between Housing Returns 1990:1-2015:1

	Belgium	Finland	France	Germany	Ireland	Italy	Netherlands	Spain

Belgium	1.000000	-	-	-	-	-	-	-
Finland	0.251948	1.000000	-	-	-	-	-	-
France	0.608360	0.599194	1.000000	-	-	-	-	-
Germany	-0.48192	-0.40174	-0.48989	1.000000	-	-	-	-
Ireland	0.437582	0.414368	0.402262	-0.376431	1.000000	-	-	-
Italy	0.576213	0.061221	0.666400	-0.504063	0.342609	1.000000	-	-
Neth.	0.332513	0.206821	0.206581	-0.327110	0.650562	0.165138	1.000000	-
Spain	0.570935	0.442953	0.799446	-0.613207	0.635165	0.761019	0.379074	1.000000

This table shows the correlations between country housing returns for the 1990:1-2015:1 sample

Table 6

Correlations between Housing Returns 1976:1-1998:4

	Belgium	Finland	France	Germany	Ireland	Italy	Netherlands	Spain

Belgium	1.000000	-	-	-	-	-	-	-
Finland	-0.28524	1.000000	-	-	-	-	-	-
France	0.581436	0.174332	1.000000	-	-	-	-	-
Germany	0.475151	-0.36628	0.307207	1.000000	-	-	-	-
Ireland	0.428779	0.146834	0.245656	0.160582	1.000000	-	-	-
Italy	-0.20619	0.143416	0.266779	0.011794	0.022355	1.000000	-	-
Neth.	0.763674	-0.18417	0.330695	0.279308	0.315207	-0.4769	1.000000	-
Spain	0.323261	0.037259	0.472761	0.043123	0.331476	0.333361	0.153125	1.000000

This table shows the correlations between country housing returns for the 1976:1-1998:4 sample

Table 7

Correlations between Housing Returns 1999:1-2015:1

	Belgium	Finland	France	Germany	Ireland	Italy	Netherlands	Spain
Belgium	1.000000	-	-	-	-	-	-	-
Finland	0.712895	1.000000	-	-	-	-	-	-
France	0.776031	0.850456	1.000000	-	-	-	-	-
Germany	-0.621318	-0.552458	-0.632573	1.000000	-	-	-	-
Ireland	0.533682	0.522089	0.587786	-0.245419	1.000000	-	-	-
Italy	0.689793	0.694086	0.824775	-0.802628	0.676986	1.000000	-	-
Neth.	0.420282	0.374150	0.497512	-0.337076	0.675999	0.694531	1.000000	-
Spain	0.662967	0.731070	0.827136	-0.699254	0.775723	0.951334	0.626219	1.000000

This table shows the correlations between country housing returns for the 1990:1-2015:1 sample

Table 8

Rank of Unconditional Bilateral Correlations, 1976:1-2015:1

Country Pair	Correlation Value	Country Pair	Correlation Value
Belg/Neth	0.69425	Germany/Neth	0.127216
France/Spain	0.57904	Germany/Ire	-0.01578
Belg/France	0.571868	Belg/Italy	-0.06338
Italy/Spain	0.45369	France/Ger	-0.08766
Ire/Spain	0.43847	Finland/Neth	-0.0913
France/Italy	0.418633	Belg/Finland	-0.12319
Ire/Neth	0.4152	Germany/Spain	-0.1397
France/Ireland	0.400188	Germany/Italy	-0.16937
Belg/Ireland	0.393475	Italy/Neth	-0.26473
Belg/Spain	0.38878	Finland/Ger	-0.3972
Finland/France	0.383533		
France/Neth	0.33082		
Belg/Ger	0.246334		
Finland/Italy	0.241462		
Neth/Spain	0.236705		
Finland/Ire	0.22977		
Ireland/Italy	0.217898		
Finland/Spain	0.1842		

Countries are ranked from highest to lowest correlation.

Table 9
Rank of Unconditional Correlations, 2008:4-2015:1

Country Pair	Correlation Value	Country Pair	Correlation Value
Italy/Neth	0.784919	France/Neth	-0.01119
Neth/Spain	0.767846	France/Ger	-0.04919
Ireland/Spain	0.735568	Belg/Spain	-0.13912
Germany/Ire	0.722519	Finland/Ire	-0.19078
Finland/France	0.604265	Finland/Neth	-0.20738
Belg/France	0.598838	France/Ire	-0.21908
Belg/Finland	0.548384	Finland/Ger	-0.29291
Belg/Italy	0.526377	Germany/Italy	-0.42508
Italy/Spain	0.507446	Belg/Ireland	-0.60035
Ger/Spain	0.373995	Belg/Ger	-0.65148
Ireland/Neth	0.370118		
Belg/Neth	0.14188		
Finland/Italy	0.094741		
Germany/Neth	0.07492		
France/Italy	0.069347		
France/Spain	0.001659		
Ireland/Italy	0.001349		
Finland/Spain	0.000863		

Countries are ranked from highest to lowest correlation.

Table 10

Unconditional Average Bilateral Correlations for Each Country

1976:1-2015:1		**1976:1-1989:4**		**1990:1-2015:1**	
Country	Correlation	Country	Correlation	Country	Correlation
France	0.3709	France	0.3833	Spain	0.425
Spain	0.3058	Belgium	0.3352	France	0.398
Belgium	0.3011	Ireland	0.3114	Ireland	0.3578
Ireland	0.297	Spain	0.2538	Belgium	0.3279
Italy	0.2141	Germany	0.2564	Italy	0.2955
Netherlands	0.2068	Netherlands	0.1832	Netherlands	0.2305
Finland	0.06103	Italy	0.04	Finland	0.2249
Germany	-0.0623	Finland	-0.1197	Germany	-0.4594

Countries are ranked from highest to lowest unconditional average correlations over their respective samples.

Table 10 (continued)

Unconditional Average Bilateral Correlations for Each Country

1976:1-1998:4		**1999:1-2015:1**		**2008:4-2015:1**	
Country	Correlation	Country	Correlation	Country	Correlation
France	0.3398	Spain	0.5535	Spain	0.450689
Belgium	0.2972	France	0.533	Italy	0.450602
Spain	0.242	Italy	0.5326	France	0.380552
Ireland	0.2358	Ireland	0.5032	Belgium	0.365281
Netherlands	0.1687	Finland	0.476	Netherlands	0.336493
Germany	0.1301	Belgium	0.4534	Ireland	0.334907
Italy	0.0135	Netherlands	0.4216	Finland	0.199398
Finland	-0.0476	Germany	-0.5558	Germany	-0.28899

Countries are ranked from highest to lowest unconditional average correlations over their respective samples.

Table 11
Economic Differences

Average Correlation	Market vs. Bank	Financial Development	Homeownership	Subsidized Renters	CA Balance
France	Ireland	Netherlands	Spain	Netherlands	Netherlands
Spain	Netherlands	Germany	Belgium	Ireland	Finland
Belgium	Germany	Ireland	Italy	Finland	Germany
Ireland	France	France	Finland	France	Belgium
Italy	Spain	Spain	Ireland	Italy	France
Neth.	Belgium	Finland	Netherlands	Spain	Ireland
Finland	Italy	Italy	France	Belgium	Italy
Germany	Finland	Belgium	Germany	Germany	Spain

Sources: Levine, R. "Bank-Based or Market-Based Financial Systems: Which is Better?", p. 45, for Market vs. bank and Financial Development, https://tradingeconomics.com/country-list/home-ownership-rate, for Homeownership, Eurostat for Subsidized Renters and the FREDS database for CA, or Current Account, Balance-this was the average current account balance as a percent of GDP 1997:1-2004:1.

Table 12
Lee-Strazicich Unit Root and Break Tests for Bilateral Correlations

	Belg/Fin	**Belg/Fran**	**Belg/Ger**	**Belg/Ir**	**Belg/It**	**Belg/Neth**	**Belg/Sp**	**Fin/Fr**
Model	Crash	Crash	Crash	Crash	Crash	Crash	Crash	Crash
T-Stat	-4.4683	-3.0816	-3.6977	-4.9968	-4.1204	-5.9891	-4.5258	-3.8625
D1 Date	1986:3	1994:2	2000:1	1993:2	1985:1	1996:3	1997:3	1987:1*
D1 Value	0.0343	-0.0038	-0.1949	-0.2043	0.2242	-0.1904	0.1186	0.3311
D1 T-Stat	0.2895	-0.03	-1.635	-1.2886	1.8671	-1.2642	0.748	2.6045
D2 Date	2005:1*	1998:1*	2008:3*	2010:4	2007:1	2007:3*	2010:4	2001:2
D2 Value	0.7855	-0.2974	-0.286	-0.1789	0.0948	0.3238	-0.1907	0.1361
D2 T-Stat	6.4433	-2.3172	-2.3826	-1.1278	0.7872	2.1757	-1.2013	1.066

The T-stat under model refers to the test statistic for the unit root test. Significant break dates are denoted with asterisks.

Table 12 (continued)

Lee-Strazicich Unit Root and Break Tests for Bilateral Correlations

	Fin/Ger	**Fin/Ir**	**Fin/It**	**Fin/Neth**	**Fin/Sp**	**Fr/Ger**	**Fr/Ir**	**Fr/It**
Model	Crash	Crash	Crash	Crash	Break	Crash	Crash	Crash
T-Stat	-5.4411	-3.4396	-3.5621	-5.9941	-2.8534	-4.0067	-4.1249	-4.0045
D1 Date	1987:1*	1985:3*	1984:3	1982:4	1986:2	1983:3	1985:4*	1990:2
D1 Value	0.3168	0.3319	-0.172	0.0621	0.0509	0.1896	0.9039	-0.2241
D1 T-Stat	2.3848	3.1092	-1.3162	0.3956	0.3926	1.1879	5.9226	-1.542
D2 Date	2011:3	1991:2	2002:3	2009:3	2000:3	1996:1	2007:3	1994:1
D2 Value	-0.1277	-0.0332	0.0617	-0.2043	0.2307	-0.2452	-0.2333	-0.0913
D2 T-Stat	-0.956	-0.2956	0.471	-1.3079	1.7669	-1.4935	-1.5101	-0.6256

The T-stat under model refers to the test statistic for the unit root test. Significant break dates are denoted with asterisks.

Table 12 (continued)

Lee-Strazicich Unit Root and Break Tests for Bilateral Correlations

	Fr/Neth	**Fr/Sp**	**Ger/Ir**	**Ger/It**	**Ger/Neth**	**Ger/Sp**	**Ir/Neth**	**Ir/Sp**
Model	Crash	Crash	Crash	Crash	Crash	Crash	Crash	Crash
T-Stat	-5.9954	-5.9693	-4.4019	-4.093	-3.0158	-3.4797	-4.0002	-5.6768
D1 Date	1994:3	1987:4	1990:1	1988:4	1993:3	1991:1	1990:2*	1987:1*
D1 Value	-0.0902	0.0306	-0.1529	-0.1497	-0.2579	0.0805	-0.5354	-0.4195
D1 T-Stat	-0.5527	0.2447	-0.9112	-0.9705	-1.7686	0.5283	-3.6954	-2.7368
D2 Date	2004:1	1995:4	1995:3	2001:2	2006:4*	2004:4*	1998:3	2007:3
D2 Value	-0.2179	-0.1985	-0.0677	-0.0305	-0.3308	-0.4713	-0.2275	-0.1661
D2 T-Stat	-1.3332	-1.5819	-0.4072	-0.1941	-2.2623	-3.2028	-1.5675	-1.066

The T-stat under model refers to the test statistic for the unit root test. Significant break dates are denoted with asterisks.

Table 12 (continued)
Lee-Strazicich Unit Root and Break Tests for Bilateral Correlations for Bilateral Correlations

	Ir/It	**It/Neth**	**It/Sp**	**Neth/Sp**
Model	Crash	Crash	Crash	Crash
T-Stat	-5.3559	-4.3226	-5.2568	-4.2581
D1 Date	2003:1	1990:2*	1986:1*	1992:1
D1 Value	0.3194	-0.3173	-0.3147	0.1977
D1 T-Stat	1.9687	-2.7063	-2.3448	1.1676
D2 Date	2007:3	2004:4	2002:1	2006:2*
D2 Value	-0.0546	0.1719	-0.1454	0.3391
D2 T-Stat	-0.3278	1.452	-1.0633	2.0781

The T-stat under model refers to the test statistic for the unit root test. Significant break dates are denoted with asterisks.

Table 13
Lee-Strazicich Unit Root and Break Tests for Unconditional Rolling Correlations

	Belg.	**Finland**	**France**	**Germ.**	**Ireland**	**Italy**	**Neth.**	**Spain**
Model	Crash	Crash	Crash	Crash	Crash	Crash	Crash	Crash
T-Stat	-4.2679	-3.6379	-4.0755	-3.8905	-5.0419	-5.4746	-4.5601	-5.9507
D1 Date	1985:4*	1987:1*	1985:4*	1985:3*	1990:1	1990:2	1991:4*	1987:4
D1 Value	0.2577	0.2	0.2687	0.2097	-0.1055	-0.1047	-0.2165	0.1242
D1 T-Stat	3.888	3.272	5.5497	2.466	-1.2998	-1.6496	-2.5523	1.8439
D2 Date	2007:2	2009:4	1990:2*	1992:2*	1999:2	1996:1	2005:2*	2006:1
D2 Value	0.1092	-0.0273	-0.1281	-0.2108	-0.0916	0.1146	-0.2165	0.0167
D2 T-Stat	1.6334	-0.4376	-2.687	-2.4797	-1.1514	-1.7757	-2.3469	0.2496

The T-stat under model refers to the test statistic for the unit root test. Significant break dates are denoted with asterisks.

Table 14
Arch Test Results

	Lags	P-Value
Belgium	2	0.0188
Finland	2	0.000
France	3	0.000
Germany	2	0.7645
Ireland	3	0.7268
Italy	2	0.000
Netherlands	4	0.000
Spain	2	0.0091

Lags refers to the number of lags used in the autoregressive model to generate residuals for the test; these lags were chosen by the SIC criterion. The p-value is from the LM ARCH test.

Table 15

DCC GARCH Conditional Mean Estimation Results

	Return Eqn.				
	Constant	AR(1)	AR(2)	AR(3)	AR(4)
BELG	0.13	1.47	-0.52		
	(0.183)	(0.00)	(0.00)		
FIN	0.105	1.64	-0.71		
	(0.49)	(0.00)	(0.00)		
FRAN	0.177	1.55	-0.49	-0.11	
	(0.048)	(0.00)	(0.04)	(0.368)	
ITALY	0.019	1.55	-0.582		
	(0.77)	(0.00)	(0.00)		
NETH	0.259	0.788	0.416	-0.059	-0.257
	(0.09)	(0.00)	(0.00)	(0.658)	(0.01)
SPAIN	0.116	1.38	-0.44		
	(0.52)	(0.00)	(0.00)		

Numbers in parentheses are p-values.

Table 16

DCC GARCH Conditional Variance Estimation Results

	Constant	a	β
Belgium	0.604	0.28	0.323
	(0.08)	(0.06)	(0.22)
Finland	1.13	0.09	0.422
	(0.00)	(0.307)	(0.00)
France	0.307	0.402	0.48
	(0.24)	(0.01)	(0.03)
Italy	0.026	0.382	0.677
	(0.15)	(0.00)	(0.00)
Neth.	0.19	0.607	0.528
	(0.09)	(0.00)	(0.00)
Spain	0.027	0.16	0.839
	(0.33)	(0.05)	(0.00)

Numbers in parentheses are p-values.

Table 17
Lee-Strazicich Unit Root and Break Tests

	Bel/Fin	**Bel/Fr**	**Bel/It**	**Bel/Neth**	**Bel/Spain**	**Fin/Fr**	**Fin/It**	**Fin/Neth**
Model	Crash	Crash	Crash	Crash	Break	Crash	Crash	Crash
T-Stat	-1.9102	-2.1068	-2.544	-2.9729	-4.9495	-3.017	-4.675	-3.722
D1 Date	1992:4	1992:1*	1986:1*	2006:1	1996:1	1981:1	2002:2	1993:2*
D1 Value	0.0769	-0.3584	0.0914	0.055	0.0587	-0.0657	0.0625	0.1302
D1 T-Stat	1.6287	-7.8244	2.4174	1.1183	1.4018	-1.408	1.5154	3.142
D2 Date	1997:2*	1997:2*	1998:1*	2009:4	1999:4*	1992:1*	2007:4*	2009:3*
D2 Value	0.1284	0.2578	0.1739	0.0729	-0.1557	0.1422	30947	-0.4801
D2 T-Stat	2.7386	5.6891	4.5926	1.3938	-3.9151	3.0419	2.2958	-11.2655
DT1 Date					1996:1*			
DT1 Value					0.0679			
DT1 T-Stat					4.0364			
DT2 Date					1999:4			
DT2 Value					-0.0066			
DT2 T-Stat					0.5554			

The T-stat under model refers to the test statistic for the unit root test. Significant break dates are denoted with asterisks.

Table 17 (continued)
Lee-Strazicich Unit Root and Break Tests

	Fin/Spain	**Fr/It**	**Fr/Neth**	**Fr/Spain**	**It/Neth**	**It/Spain**	**Neth/Spain**
Model	Crash	Crash	Crash	Crash	Crash	Break	Break
T-Stat	-1.8212	-2.5759	-3.45	-1.9106	-1.837	-3.988	-4.8769
D1 Date	1982:4*	1997:1*	1992:1*	1983:1	1993:2*	1984:1	1985:4
D1 Value	-0.1033	0.1088	-0.3413	-0.0816	-0.1526	-0.0034	0.0598
D1 T-Stat	-2.6863	2.7781	-6.5951	-1.7175	-4.1527	-0.1037	1.2419
D2 Date	1992:1*	2001:4	1996:1*	1997:2*	2009:4*	2008:1	1997:4
D2 Value	0.1744	0.0553	0.0974	0.1319	0.11	-0.0101	-0.0656
D2 T-Stat	4.598	1.4349	1.9633	2.7427	2.7545	-0.303	-1.3496
DT1 Date						1984:1*	1985:4*
DT1 Value						0.0207	0.0401
DT1 T-Stat						2.869	3.0563
DT2 Date						2008:1	1997:4*
DT2 Value						0.0163	0.0519
DT2 T-Stat						1.6212	4.0403

The T-stat under model refers to the test statistic for the unit root test. Significant break dates are denoted with asterisks.

Table 18
Lee-Strazicich Unit Root and Break Tests for Similarity

	Total	Belgium	Finland	France	Germany
Model	Crash	Crash	Crash	Break	Break
T-Stat	-6.5041	-6.1332	-6.8941	-6.2729	-6.8008
D1 Date	1999:4*	1982:1*	1986:1*	1996:1*	1990:2
D1 Value	0.1927	0.5771	-0.5196	-0.1805	0.0861
D1 T-Stat	2.5496	2.709	-1.2923	-4.645	0.3674
D2 Date	2010:1	2006:2*	1993:1	2009:4*	1994:4
D2 Value	-0.0067	0.4313	0.4507	-1.3146	-0.2754
D2 T-Stat	-0.0877	2.0178	1.1202	-4.6649	-1.2126
DT1 Date				1996:1*	1990:2
DT1 Value				0.3127	0.0933
DT1 T-Stat				4.6965	1.4942
DT2 Date				2009:4*	1994:4*
DT2 Value				0.3866	-0.1347
DT2 T-Stat				4.8745	-2.182

The T-stat under model refers to the test statistic for the unit root test. Significant break dates are denoted with asterisks.

Table 18 (continued)
Lee-Strazicich Unit Root and Break Tests for Similarity

	Ireland	**Italy**	**Netherlands**	**Spain**
Model	Break	Break	Crash	Crash
T-Stat	-6.7668	-6.2576	-4.6884	-6.9395
D1 Date	1991:4	1979:4	1980:4*	1980:2
D1 Value	0.0252	-0.183	0.9025	-0.6235
D1 T-Stat	0.0612	-0.722	2.124	-1.3691
D2 Date	1997:2	1993:4	1995:2*	1987:3
D2 Value	-0.2484	-0.3368	-1.2966	0.3723
D2 T-Stat	-0.5914	-1.337	-3.0832	0.8167
DT1 Date	1991:4	1979:4*		1980:2
DT1 Value	-0.0051	-0.1853		-0.1036
DT1 T-Stat	-0.0507	-2.4108		-0.7606
DT2 Date	1997:2	1993:4*		1987:3*
DT2 Value	0.5927	-0.1045		0.6992
DT2 T-Stat	4.7211	-2.279		5.3063

The T-stat under model refers to the test statistic for the unit root test. Significant break dates are denoted with asterisks.

Figure 1
Belgium Rolling Unconditional Average Correlation

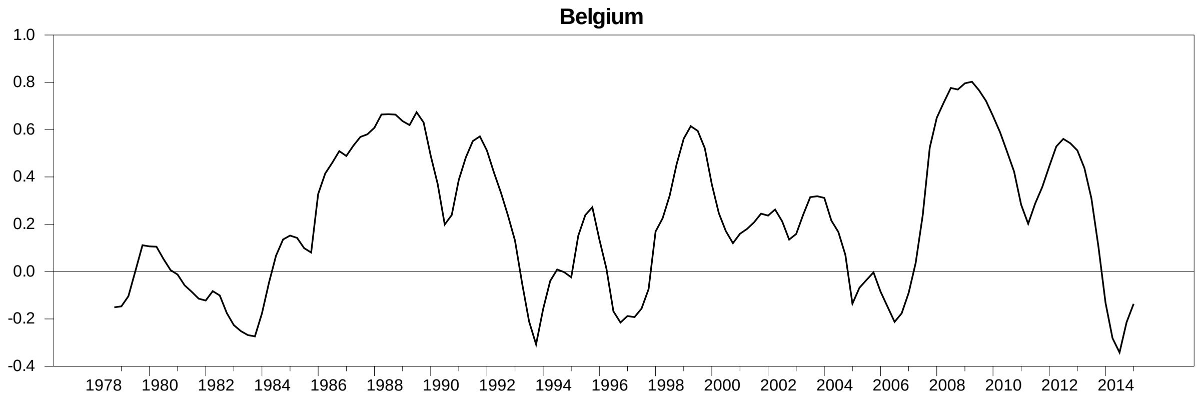

Figure 2
Finland Rolling Unconditional Average Correlation

Figure 3
France Rolling Unconditional Average Correlation

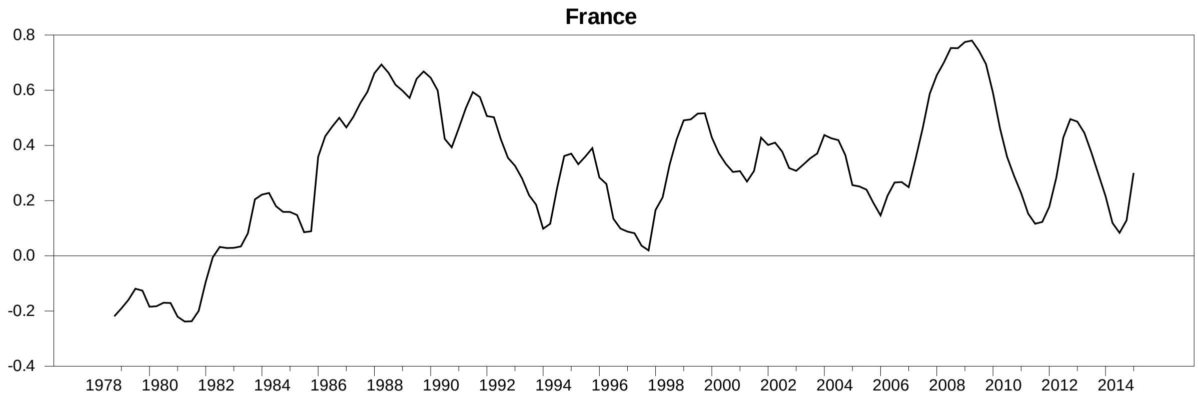

Figure 4
Germany Rolling Unconditional Average Correlation

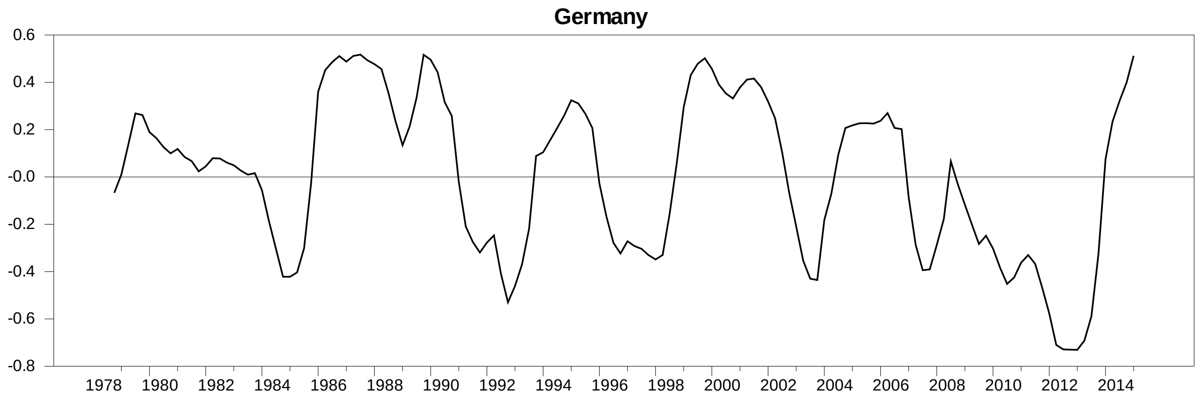

Figure 5
Ireland Rolling Unconditional Average Correlation

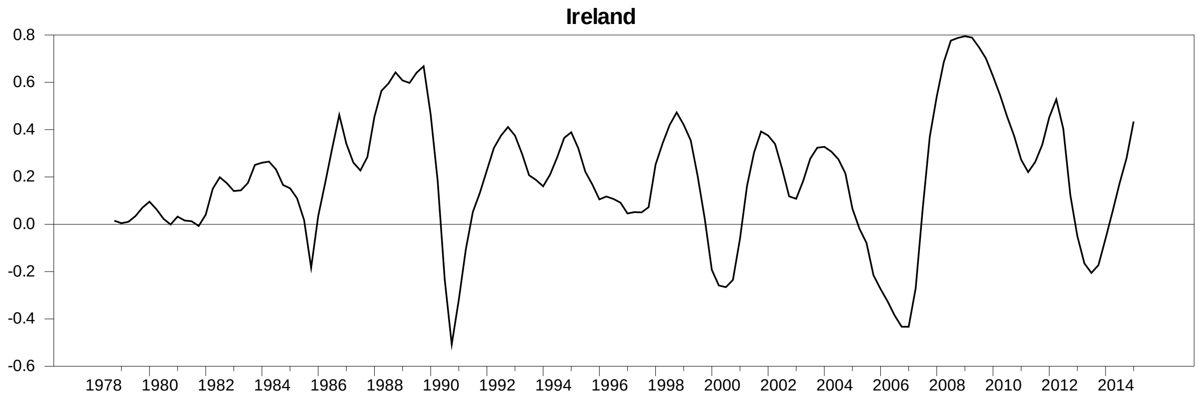

Figure 6
Italy Rolling Unconditional Average Correlation

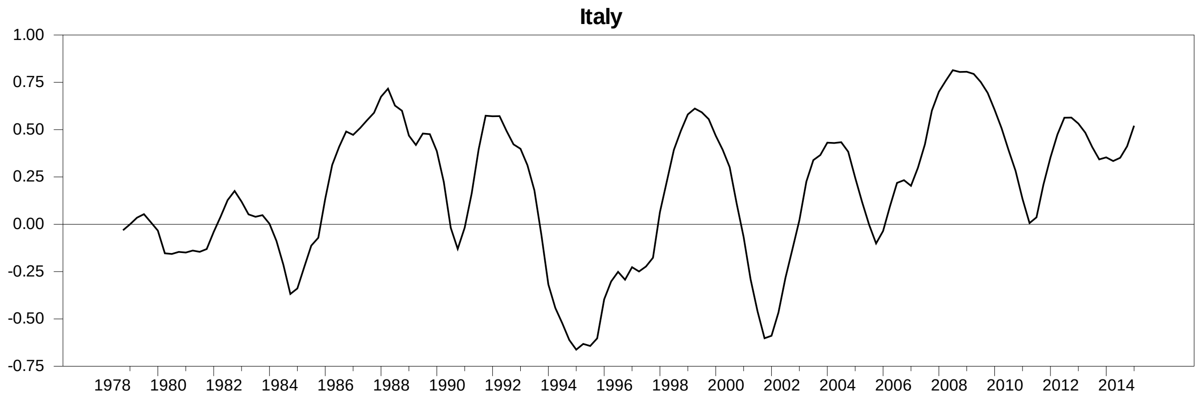

Figure 7
Netherlands Rolling Unconditional Average Correlation

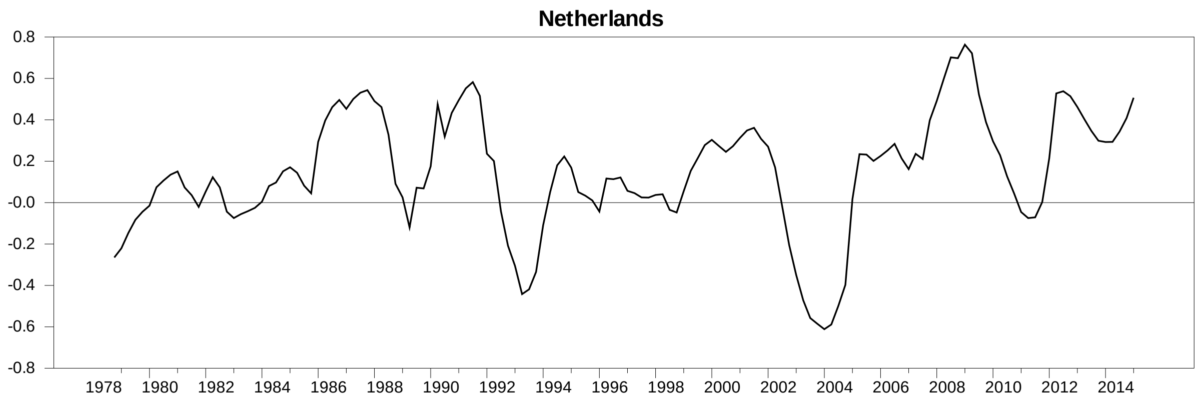

Figure 8
Spain Rolling Unconditional Average Correlation

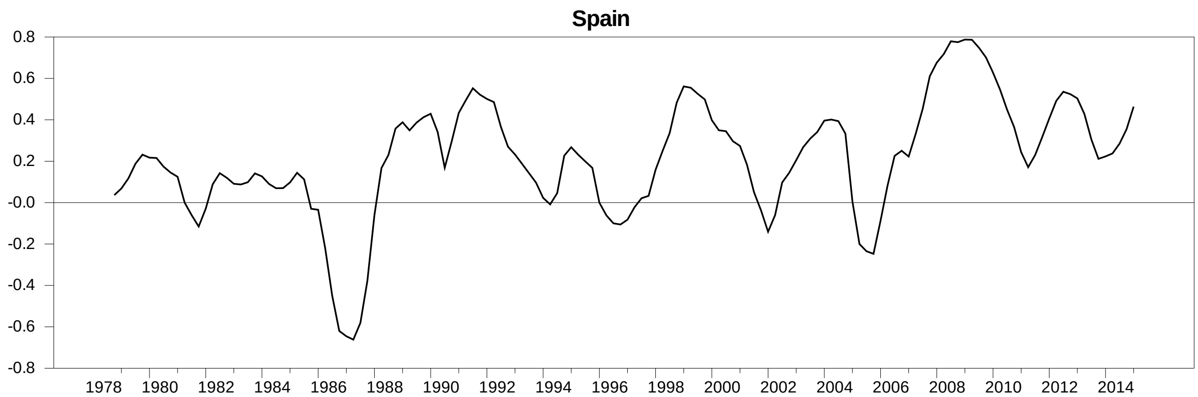

Figure 9
Overall Similarity

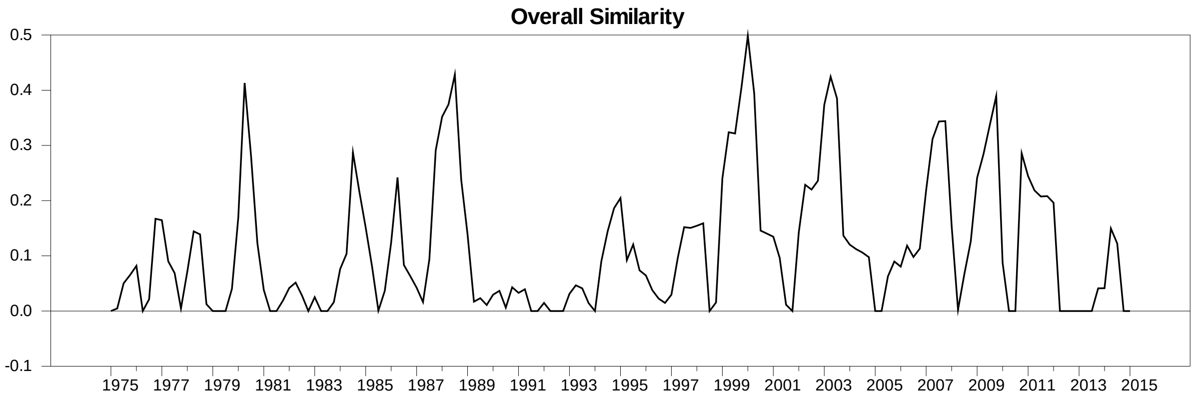

Figure 10
Belgium Similarity

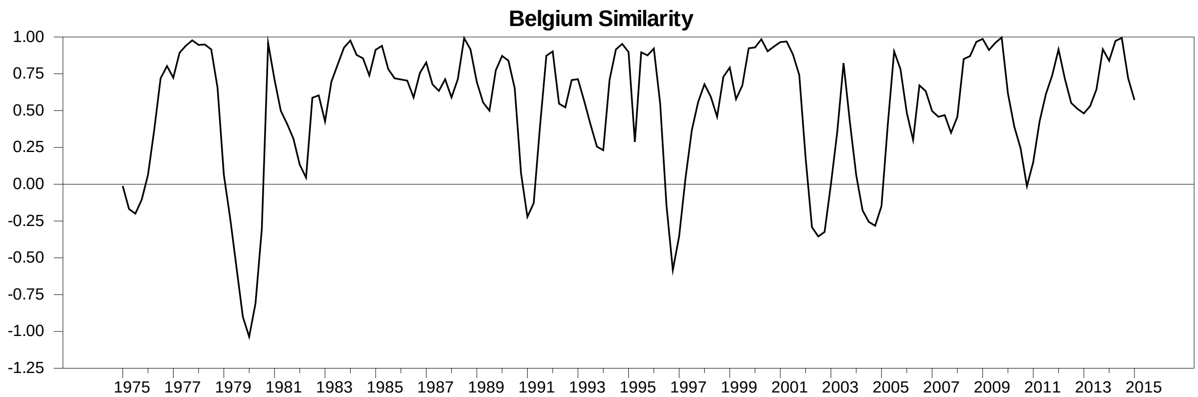

Figure 11
Finland Similarity

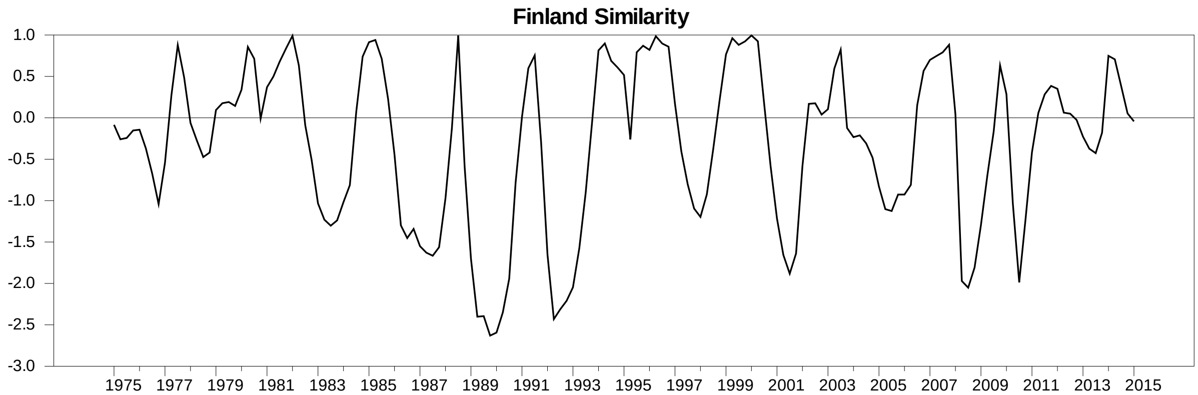

Figure 12
France Similarity

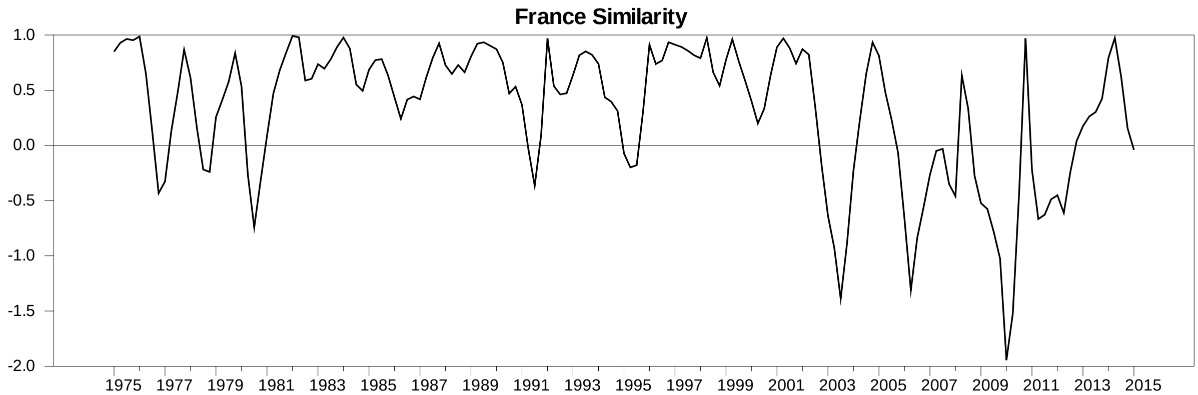

Figure 13
Germany Similarity

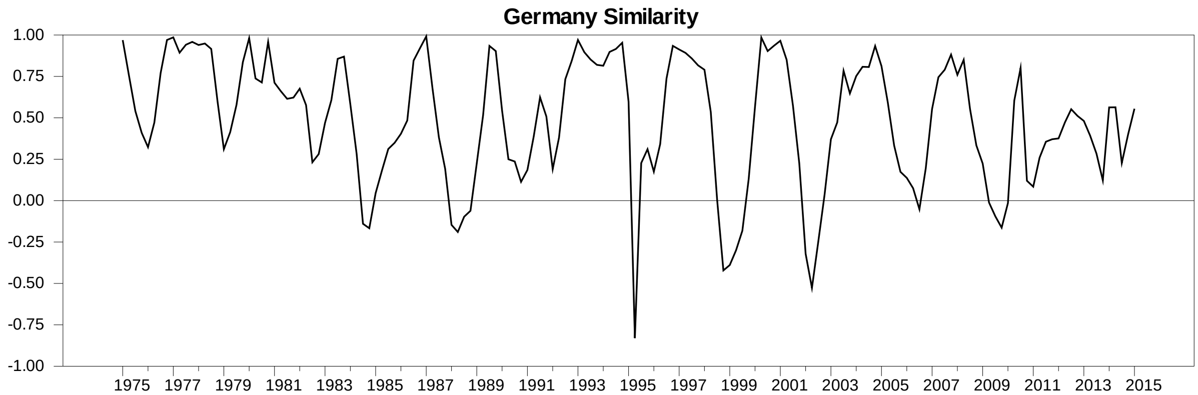

Figure 14
Ireland Similarity

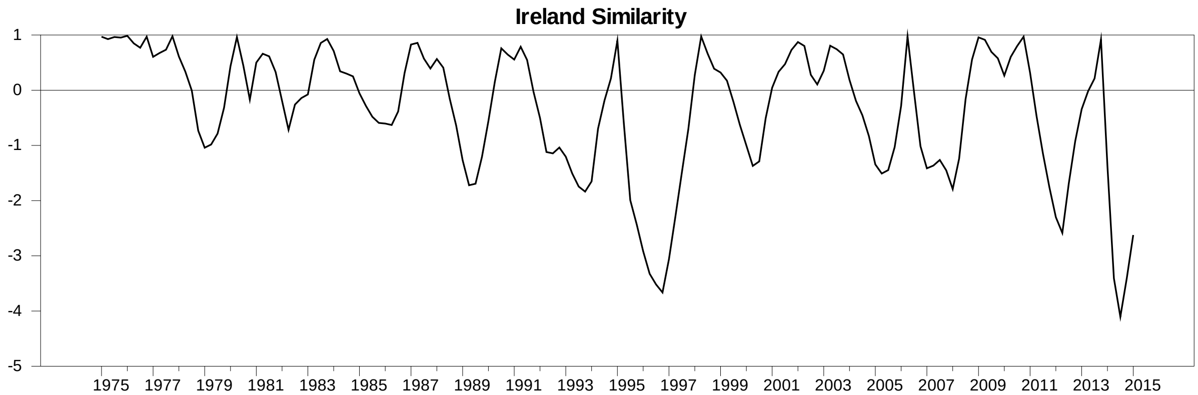

Figure 15
Italy Similarity

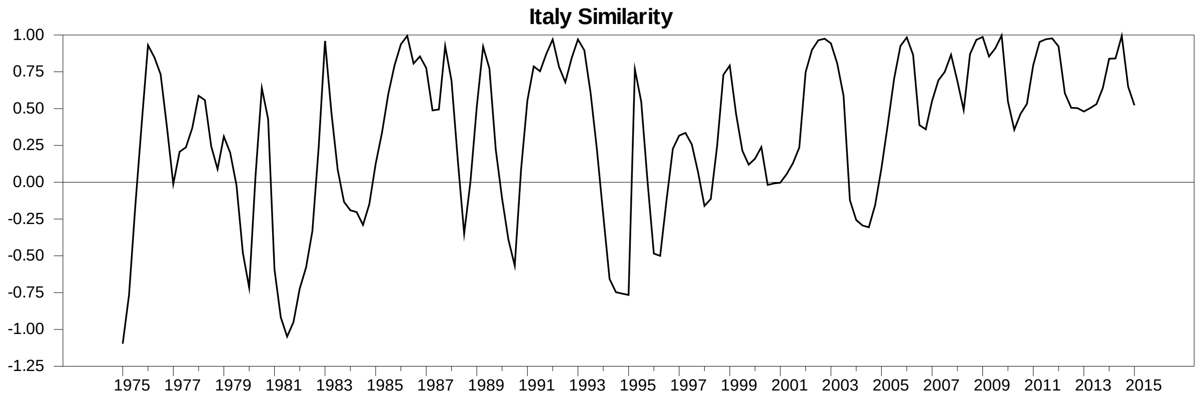

Figure 16
Netherlands Similarity

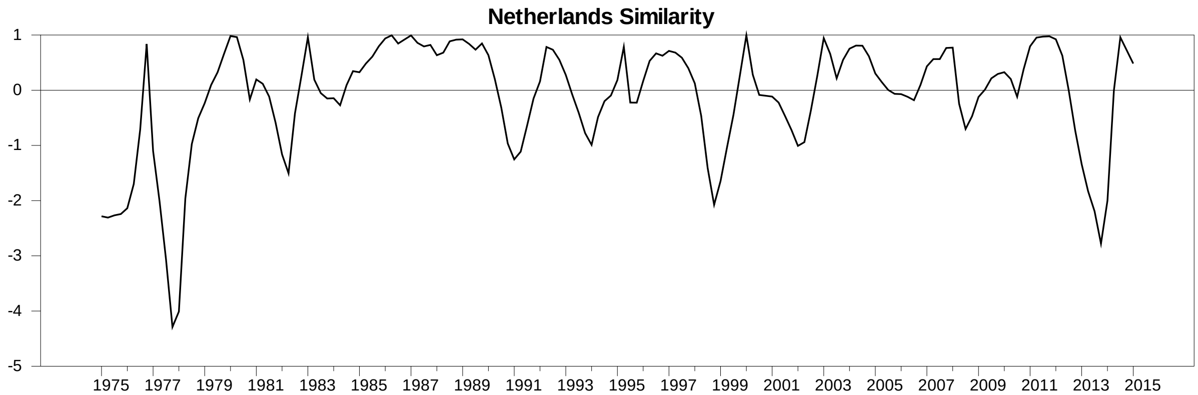

Figure 17
Spain Similarity