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International Migration to Germany: Estimation of a Time-Series Model and Inference in Panel Cointegration.*

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Abstract

In this paper we study the determinants of international migration to Germany, 1967-2000. The empirical literature on macro-economic migration functions usually explains migration flows by a set of explanatory variables such as the income differential, employment rates, and migrations stocks as in Hatton (1995), for example. Since macroeconomic variables are widely acknowledged as non-stationary, the standard model in the migration literature can only meet the requirements of modern non-stationary time-series econometrics if migrations flows and the explanatory variables are integrated of the same order and if these variables form a cointegrated set. In order to prove whether the standard specification is compatible with our data, we use the univariate Augmented Dickey-Fuller test as well as its panel data version, developed in Im, Pesaran, and Shin (2003), to test for unit roots in the time series. The tests demonstrate that migration rates are stationary, while the remaining explanatory variables follow I(1) processes. Consequently, we suggest an alternative specification of the long-run migration function with migration stocks as the dependent variable. For this specification, we find that all variables are I(1) processes, and that the null of no cointegration can be decisively rejected by applying the panel cointegration test of Pedroni (1999). The parameter inference in the cointegrating regressions is conducted using the method of canonical cointegrating regressions of Park (1992). Our empirical findings generally agree with predictions of migration theory.

Keywords: Migration, unit roots, panel cointegration.

JEL Classification Codes: F22, C22, C23

1 Introduction

In this paper, we analyze the macroeconomic determinants of international migration into Germany from a panel of source countries from the European Union (EU). Our focus on Germany is not purely arbitrary. Being one of the largest economies in the world, Germany serves as an important migration destination country: it is home to around 40% of the EU citizens who live in other EU countries (Eurostat, 2000). Moreover, the analysis of the factors determining migration into Germany is greatly facilitated by the fact that the country is one of the few in the EU that consistently reports international migration flows and stocks on an annual basis by country of origin over a long span of time.

In general, the analysis of the economic determinants of international migration is hampered by the fact that institutional restrictions naturally hinder the effectiveness of the economic forces that have an impact on the migration decisions of economic agents. The European Union (EU) forms an interesting exception in this respect, since it has eliminated the legal and administrative barriers to the movements of workers and other persons within the community to a large extent. In 1968, free movement was guaranteed to the citizens of the six founding members of what was then the European Economic Community (EEC) with a joint population of 185 millions, and has since been extended to the 18 members of the EU and the European Economic Area (EEA) with a joint population of 375 million. Thus, the EU and the EEA form a natural laboratory in which to study the economic determinants of international migration.

In this paper, we consider migration to Germany from the five other founding members of the EU, where free movement was introduced in 1968, and three further countries, where free movement was introduced when they joined the EU in 1973. Hence it is assumed that for the present subsample of the migrant population, the decisions to emigrate to Germany were mostly voluntary and the emigration occurred without targeted government intervention of an either encouraging or discouraging nature over the relevant course of time.

In the empirical literature on macro migration functions, usually gross or net migration flows are explained by a set of variables. The choice of the explanatory economic variables is motivated by microeconomic theories of migration, where individuals form expectations on utility differences in the respective locations, which are determined by the income differentials between home and host countries as well as by the variables that reflect labor market conditions in the respective locations, see e.g. Hicks (1932), Todaro (1969), and Harris and Todaro (1970). Moreover, many models include the existing migrant stock as a proxy for 'social network' effects which are expected to increase the propensity to migrate by alleviating the adaptation costs in the host country, see Stark (1991) and Epstein and Hillman (1998).

Hatton (1995) applies the recent advances in non-stationary time series econometrics in estimating macro migration functions. Largely in line with the literature, his migration function relates gross (net) migration rates to several explanatory economic variables as well as to the already existing migrant stock in the destination country. In contrast to the existing migration literature, he addresses the statistical properties of the data by testing for the order of integration of the economic variables. The finding that the relevant variables are found to be I(1) subsequently leads to the statistically supported hypothesis that these variables form a cointegration set. Thus, the resulting migration function of Hatton (1995) represents the long-run equilibrium relation between the migration flows on the one hand and the explanatory economic variables on the other.

Observe that Hatton (1995) model implies that both migration flows and migration stocks variables are I(1). It is however rather unlikely that both these variables are integrated of the same order, since migration flows essentially are the first difference of migration stocks. Hence, we expect that if the migration stock variables are I(1), then the corresponding flow variables should be I(0), according to the definition. We suggest therefore in this paper an alternative specification of the long-run migration function, where migration stocks are explained by income differential and employment variables. In this form of migration function the problematic issue of the same order of integration of migration flows and stocks variables is no longer present.

The hypothesis that a long-run equilibrium between migration stocks and the explanatory variables exists can be motivated by the assumption that individuals differ with respect to their preferences and human capital characteristics, which in turn affect benefits and costs of migration. Consequently, for a given difference in expected income, stock of migrants will eventually achieve an equilibrium, where benefits and costs of migration are equalized for the marginal migrant.

Taking these considerations as a starting point, we formulate the objectives of this paper as follows: Firstly, we shall determine whether the properties of the data at hand can be reconciled either with the migration flow or the migration stock equation. To this end, we shall determine whether the order of integration of migration flows or migration stocks is compatible with that of the explanatory variables using both single time series and panel data unit root tests. Secondly, we shall test whether the hypothesis of a dynamic equilibrium relationship between migration rates or, alternatively, migration stocks on the one hand, and macroeconomic variables such as per capita income differential and employment rates on the other, are supported by our data. Thirdly, after testing for cointegration, we estimate the heterogeneous cointegration vectors in order to make inferences on the parameter estimates of the long-run relations between migration and the relevant explanatory economic variables.

The remainder of the paper is organized as follows. Section 2 discusses the theoretical considerations behind the macro models of migration. Section 3 describes the database used. Section 4 presents the empirical results. The final section concludes.

2 Theoretical Considerations

A standard specification of the long-run migration function in the empirical literature has the following form (see e.g. Hatton, 1995):

$$\ln(m_{it}) = a_{0i} + a_{1i} \ln\left(\frac{w_t^*}{w_{it}}\right) + a_{2i}\ln(e_t^*) + a_{3i}\ln(e_{it}) + a_{4i} mst_{it} + a_{5i} time + u_{it}, \quad (1)$$

where $i = 1, ..., N$ and $t = 1, ..., T$ are the time and country indices, m_{it} denotes the gross migration rate as percentage of the whole home population of country i, w_t^* the wage rate in the host country, w_{it} the wage rate in the home country i, e_t^* the employment rate in the host country, e_{it} the employment rate in the home country i, mst_{it} the migrant stock variables measured as a percentage of the whole home population of country i, and $time$ stands for a linear deterministic time trend. Observe that the specification of the migration function in double logarithmic form is common in the literature (see e.g. Faini and Venturini, 1996; Hille and Straubhaar, 2001), but only one among many different functional forms that can be developed. Based on other assumptions about the utility function, Hatton (1995) also conceives a semi-logarithmic specification.

This parsimonious specification of the migration function has a long tradition in the literature. The choice of economic variables is primarily based on the classical contributions of Ravenstein (1889), Hicks (1932), Sjaastad (1962) as well as Todaro (1969) and Harris and Todaro (1970). The former studies suggest that differences in wages and other sources of income between the host and the source countries could be regarded as a primary determinant of migration decision, while the latter two papers introduce the role of labor markets in the decision-making process.[1]

More specifically, Hatton (1995) presents a migration model that is derived from the following theoretical assumptions. The utility of risk-averse individuals is determined by expectations on income levels in the respective locations. Utility is convex in the income differential. Uncertain employment opportunities affect the expectations on income levels in the foreign and home countries. Migration networks alleviate the costs of adapting to an unfamiliar environment, such that the costs from migration are expected to decline with the stock of migrants already existing in the host country. The time trend serves as a proxy for the variation in the costs of migration, which are expected to fall over time in the course of decreasing

[1] Microeconomic models of the migration decision have gone far beyond these classical contributions. *Inter alia*, these models analyze the role of financial constraints in the absence of perfect capital markets, see Stark (1991), the impact of uncertainty about future wage and employment conditions on the migration decision in the presence of fixed migration costs, see Burda (1995), the choice of the optimal length of stay in a foreign country if migration is temporary, see Djajic and Milbourne (1988). However, few of these theoretical contributions have developed macro migration functions whose parameters can be estimated empirically.

transport and communication costs.

Thus, it is expected that the income differential and the host country employment rate have positive coefficients, while the coefficient of the employment rate in the home country has a negative sign. Moreover, positive signs are expected for the coefficients of the stock of migrants and of a deterministic time trend. In addition, the model predicts that the coefficient for the employment variables is larger than that for the income variables. Finally, since employment opportunities of migrants in host countries are below those from natives, it is expected that the coefficient for the employment rate in the host country is larger than that in the source country.

In this paper we want to highlight a key assumption of the standard model in equation (1), which is largely ignored in the literature. The model presumes that a log-linear relationship between the migration flow and the economic variables exists in the long-run equilibrium. This implies that migration ceases not before (expected) income levels between the host and the source country, as determined by the wage and the employment variables on the right hand side of equation (1), have converged to a certain threshold level, which is determined by the fixed and variable costs of migration. In case of persistent differences in (expected) income levels, either the total population will eventually migrate or migration will not happen at all from the beginning.

Note that this is a consequence of deriving macro migration functions from the concept of a representative agent, i.e. of assuming that individuals are *homogenous*. If we assume that individuals are *heterogenous*, i.e. that they differ with regard to their preferences and their human capital characteristics, which in turn determine the benefits and costs of migration, (net) migration ceases when the benefits from migration equals its costs for a marginal migrant. Thus, we achieve an equilibrium between migration *stocks* instead of migration *flows* and the difference in (expected) income levels in the respective locations. As an alternative to the standard model in equation (1), we conceive therefore the following specification for the long-run migration function:

$$\ln(mst_{it}) = b_{0i} + b_{1i} \ln\left(\frac{w_t^*}{w_{it}}\right) + b_{2i} \ln(e_t^*) + b_{3i} \ln(e_{it}) + b_{4i} time + v_{it}. \tag{2}$$

The estimation of the migration functions in equations (1) and (2) can be affected by spurious correlation problem, if the regressions involve integrated of order one, I(1), variables (see the seminal paper by Granger and Newbold, 1974). The notable exception is the situation when I(1) dependent and explanatory variables form a cointegration set, see Engle and Granger (1987). From this point of view, equation (1) should represent a cointegrating equation. Indeed, the Hatton (1995) paper provides the supporting evidence of this proposition.

However, the possible empirical caveat with the specification of equation (1) is that the underlying assumption that all the variables should be I(1) might not be supported by other data sets. While there is general agreement that macroeconomic variables such as income levels and employment rates are rather well represented as I(1) processes, there still is limited evidence on the time series properties of the migration flows and corresponding migrant stock variables. Particularly puzzling is the fact that both the migration flow and the migration stock variable are included in equation (1). Since migration flows can be conceived as (almost) the first difference of migration stocks, they can hardly be I(1) variables if migration stocks are supposed to be I(1) variables as well. Thus, we expect the migration flow variable can be better approximated by an I(0) process if we find that migration stocks are I(1). In this case we suggest to apply our alternative specification in equation (2) for estimating the long-run migration function.

Thus, the purpose of the empirical section below is to test whether the specification of the migration function of Hatton (1995), which worked well in describing UK-US migration during the time period 1870-1913, is compatible with our data at hand, or whether we should conceive an alternative specification along the lines of equation (2). Moreover, we estimate the long-run migration function in order to test whether the signs and the relative magnitudes of the coefficient estimates support our theoretical expectations.

3 Data

Our sample comprises the migration data from eight European source countries to Germany: Belgium, Denmark, France, Italy, Ireland, Luxembourg, the Netherlands, and the United Kingdom. The data on migration stocks and flows come from the Federal Statistical Office (*Statistisches Bundesamt*) in Germany, see Statistisches Bundesamt (2001). For the stock of migrants, we used the foreign residents as reported by the Central Register of Foreigners (*Ausländerzentralregister*) as a variable.[2] This data is available from 1967 to 2000. The stock of foreign residents is reported on December 31 (in some early years on September 30). The number of foreign residents is slightly overstated by the Central Register of Foreigners, since return migration is not completely registered by the municipalities. Consequently, the figures for the stock of foreign residents has been revised two times in the wake of the population censuses of 1972 and 1987. Moreover, after German unification, complete figures for Western Germany are no longer available. Since the number of foreigners in Eastern Germany has been fairly low, this does not affect the total figures much. The data on migration flows stem again from the Central Register of Foreigners. The migration stock and flow variables are calculated as shares of the corresponding home population. Population figures are depicted from the World Bank World Development Indicators 2000

[2]Note that all residents in Germany are obliged to register themselves at the place of residence. The figures from the central register of foreigners are based on the reports of the municipalities.

and OECD sources.

As a proxy for wages and other incomes, we use per capita GDP as reported by the Main Economic Indicators of the OECD. The employment rate is defined as one minus the unemployment rate. Unemployment rates have been taken again from the OECD Main Economic Indicators, where this was not available, data from national statistical offices was used. The ILO definition has been used for all unemployment rates.

4 Empirical results

As noted in Hatton (1995), the long-run migration model (1) has performed well on his data. All variables appear to be I(1), and, moreover, they represent a cointegration set. Therefore the first thing to check, if we want to apply his model specification to our data, is to infer the order of integration of our country-specific variables. To this end, we use the Augmented Dickey-Fuller (ADF) test for the individual time series and the panel unit root test, suggested in Im, Pesaran, and Shin (2003) (IPS-test).

In the econometrics literature, the use of panel unit root tests has been justified by their significantly greater power when compared to univariate tests (see e.g. Levin, Lin, and Chu, 2002). Indeed, the results of panel unit root tests applied to a various economic time series, such as nominal interest rates in Wu and Zhang (1996), inflation rates in Culver and Pappell (1999), and real exchange rates in Wu (1996) and Papell (1997), *inter alia*, often yield opposite results to those of the univariate unit root rests. In particular, Wu (1996) observes that he never was able to reject the unit root hypothesis in real exchange rates using univariate unit root tests (e.g. Augmented Dickey-Fuller and Phillips-Perron tests) at the conventional significance levels, whereas the panel unit root tests reject the null hypothesis of unit root even at the 1% significance level. Hence it is useful to check whether the univariate and panel data unit root tests yield consistent results when applied to our data.

Insert Table 1 about here.

Table 1 reports the results of the ADF and IPS unit root tests performed on the host and home country-specific economic variables. The choice of the deterministic terms in the ADF regressions, i.e. the inclusion of an intercept or an intercept and a linear trend, was based on the following considerations for each variable separately: for those variables that clearly exhibit a trending behavior, we included a linear trend in order to increase power against the possible trend-stationarity of the variable. These variables are the migrant stock and employment rate in Germany. For other variables, where no such clear-cut distinction can be made, we report the results of tests with both sets of deterministic components.

As expected, the results of the unit root tests suggest that the null hypothesis that our macroeconomic explanatory variables, i.e. the relative income ratio and the employment rates, follow I(1) processes, cannot be rejected. We also cannot reject the null hypothesis that the migrant stock variables are I(1) processes as well. What is striking is that even so the univariate ADF test results applied to gross migration rates suggest that in some cases the null hypothesis of a unit root cannot be rejected, the more powerful panel data unit root tests yield the opposite conclusion. Thus, we can reject the null unit root hypothesis for gross migration rates. Hence, based on the unit root results, we conclude that our data seem to be incompatible with the specification of the long-run migration function suggested in Hatton (1995). The regression equation is unbalanced as the chosen dependent variable (gross migration rates), which has been found to be I(0), is being explained by the non-stationary I(1) variables.

Thus, in order to reconcile the features of our data with the theoretical considerations, we utilize the long-run migration function specified in equation (2) above. According to the unit root test results, all the variables of regression (2) seem to be I(1) and they should hypothetically form a cointegration set in order to avoid the 'spurious' regression effects. Under the assumption of cointegration, the remainder term v_{it} is assumed to be an I(0) variable.

We use two specifications of this cointegrating relation: without including a linear deterministic trend, i.e. $b_1 = 0$, and with a linear trend, i.e. $b_1 \neq 0$. In economic terms, the presence of a linear trend in our regression helps to account for the constant growth rate in the migration stock that has been caused by other factors than income differential and employment conditions. These socioeconomic factors that are not modelled explicitly reflect the decreasing moving costs over time caused *inter alia* by the increasing integration of the European economy.

Our next task comprises testing the hypothesis that the variables entering equation (2) are cointegrated. We report two sets of the results of cointegration tests. The first set comprises the results of the two-step Engle-Granger cointegration procedure performed for the variables of every country. The second set of results comprises the panel cointegration group $t-$test statistics of Pedroni (1999) which aggregates the test statistics of the Engle-Granger procedure reported for every country in our panel.

Table 2 contains the results of the time series and panel cointegration tests of the relation of interest given in equation (2) for both specifications with and without a linear deterministic trend. For the specification without trend, we reject the null hypothesis of no cointegration for the four countries out of the eight considered at the 10% significance level and only once at the 5% significance level. Here we see that, the inclusion of the trend in the model resulted in the rejection of the null hypothesis in four cases at the 10% level and for three countries at the 5% significance level. The panel cointegration test of Pedroni (1999) rejects however the null hypothesis of no cointegration for both model specifications at the 5% significance level.

Insert Table 2 about here.

The results of the cointegration tests suggest that every group of country-specific variables form a cointegrated set. Given this, we intend to estimate and to make an inference on parameter estimates of the heterogenous country-specific cointegrating relations. Unfortunately, the distribution of the OLS estimator of the cointegrating regression (that we employ in the two-step Engle-Granger method) is unknown for the general case due to the static nature of the cointegrating regression and unaddressed 'endogeneity bias', see e.g. Patterson (2000) for details. Therefore we are unable to make a valid inference on the parameters of interest. Hence, we estimate the cointegrating vectors by the method of canonical cointegrating regression (CCR, in short) of Park (1992), which is the nonparametric method of transforming the variables in a way that eliminates the unfortunate properties of the OLS estimator. The CCR estimator has a standard normal asymptotic distribution and therefore makes the standard parameter inference possible.

Insert Table 3 about here.

The estimation results of cointegrating vectors are presented in Table 3, where the left(right) panel contains the slope parameter estimates of cointegrating equations when a linear deterministic trend has been included(omitted) in equation (2). Recall that according to the theoretical considerations discussed in Section 2, we expect positive signs for the difference in per capita GDP levels as well as for the employment rate in the host country. Note also that a negative sign is expected for home employment rates since they increase employment opportunities in the source country. However, several empirical studies find that home employment opportunities have actually increased migration (see Greenwood, 1975, for a review). One possible explanation for this phenomenon is that favorable employment opportunities at home might help to lift liquidity constraints that a potential migrant experiences there, and, in doing so, might prove to be a factor encouraging migration.

When comparing the estimation results for both model specifications, i.e. with and without trend, we notice that the regressions with a time trend yield more plausible results. In the model with the linear deterministic trend we find that five coefficients for the income ratio are significant at the 5% level, four of which are positive. Interestingly enough, with the exception of UK, all coefficients of the German employment variable have the expected positive sign. Five of the positive coefficients are significant at the 5% level. Moreover, the coefficients of the employment variable are larger than the coefficients of the income variables in all cases where both variables have the expected positive sign. This is in line with the expectations which follow from the theoretical model. Finally, the home employment rate coefficient is found to be significant in five cases at the 10% level, out of which three estimates have a positive, and the remaining two - a negative sign. Note that the inclusion of a linear trend has wiped out the implausibly high values of the long-run elasticities of the migrant stock with respect to the host country employment rate in case of Ireland and UK.

In the cointegrating regression without trend, the income ratio variable is significant for the four countries at the 10% level, and in three cases it has the expected positive sign. The German employment rate variable is found to be significant in six cases at the 10% level, and in only two cases does it have the expected positive sign. Finally, we find that the home employment rate variable is significant in only three cases at the 10% level and has a negative sign in all significant cases.

5 Conclusions

In this paper we analyzed the determinants of international migration from eight European source countries to Germany in the period 1967 to 2000. We chose these countries because they did not restrict emigration for most of the time period. This enables us to draw interesting conclusions on the economic forces underlying international migration.

The standard approach in the empirical literature presumes the existence of an equilibrium relationship between the migration rate and explanatory variables such as the difference in income and employment opportunities between the host and the home country; see Hatton (1995) as an example. In terms of modern non-stationary time series econometrics, the migration function has to be modelled within a cointegration framework, which demands that both the migration rate and all the explanatory variables are non-stationary I(1) variables, and that a long-run equilibrium between the economic variables exists, i.e. that the variables in question are cointegrated. The statistical tests in the Hatton (1995) paper indeed suggest that for UK-US migration between 1870 and 1913 the long-run migration function meets all the necessary requirements. However, following a well-established tradition in the empirical literature, the Hatton model includes migration flows and stocks in one equation, which theoretically cannot form a cointegrated set if migration stocks are I(1) variables on the one hand, and migration flows are first differences of migration stocks, i.e. I(0) variables, on the other. We considered therefore an alternative specification, which avoids this problem by assuming that an equilibrium relationship between migrant stocks and the explanatory variables exists.

One the one hand, we find analogously to the Hatton (1995) paper that the difference in per capita GDP levels as well as the employment rates could be approximated by non-stationary stochastic processes. On the other hand, however, the hypothesis of a unit-root is rejected for the migration rate in our data. This latter finding contrasts clearly the results of Hatton (1995). This implies that the flow model is incompatible with our data, as the stationary migration rate and non-stationary variables such as the difference in per capita GDP and employment rates cannot be cointegrated, and thus cannot form a long-run dynamic equilibrium relationship. However, we find evidence supporting the existence of cointegration between the migration stocks and the key explanatory variables such as the difference in per capita incomes, and the host and source countries employment rates.

Our results generally support the theoretical hypothesis that a positive difference in per capita income between host and home country increases the long-run stock of migrants. However, the impact of employment rates in the home countries on migration has turned out to be ambiguous in our sample.

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Table 1: Results of the ADF and IPS tests

	$\ln(mst_{it})$		$\ln(m_{it})$				$\ln(e_{it})$				$\ln(e_t^*)$		$\ln(w_t^*/w_{it})$			
	with trend		without trend		with trend		with trend		without trend		with trend		with trend		without trend	
Belgium	-2.36	0	-2.95*	1	-2.93	1	-0.96	1	-1.65	1	-3.02	1	-1.42	0	-1.37	0
France	-2.82	0	-3.46**	1	-3.33*	1	-0.80	1	-1.44	1	-3.02	1	-3.54*	0	-3.10	0
Italy	-2.03	4	-2.27	1	-3.07	1	-2.89	1	-1.26	1	-3.02	1	-1.52	0	-1.92	0
Luxembourg	-3.58*	0	-2.33	1	-2.04	1	-4.38***	2	-0.75	3	-3.02	1	1.18	0	-0.22	1
Netherlands	-2.93	0	-2.55	1	-2.67	1	-1.11	1	-1.90	1	-3.02	1	-1.50	0	-1.35	0
Denmark	-3.92**	4	-1.17	3	-2.20	1	-1.22	1	-1.74	1	-3.02	1	1.09	5	-0.76	5
Ireland	-0.86	0	.NaN	.NaN	.NaN	.NaN	-0.44	1	-1.45	1	-3.02	1	1.85	0	-1.77	4
UK	-1.26	0	-3.09**	3	-3.65**	1	-0.67	2	-1.65	2	-3.02	1	-1.85	3	-2.80	3
IPS-test	-1.12		-2.99***		-2.06**		1.95		0.08		.NaN		2.44		1.49	

Table reports the results of the the Augmented Dickey-Fuller (ADF) and Im, Pesaran, and Shin (IPS) test statistics with the selected augmentation lag length according to Campbell and Perron (1991). The critical values of the ADF test are reported after Hamilton (1994) for the sample size of 25 observations. They are -3.75, -3.00, and -2.63 for the corresponding 1%, 5%, and 10% significance levels when only an intercept is included in the ADF test regression and -4.38, -3.60, and -3.24 - when an intercept as well as a linear time trend are included, see Hamilton (1994). The signs ***, **, and * denote the significant test statistics at the 1%, 5%, and 10% levels, respectively.

The IPS test statistics is reported in the last line of the table. The IPS test statistic has the standard normal asymptotic distribution. The one-sided critical values are -2.32, -1.64, and -1.28 for the respective 1%, 5%, and 10% significance levels.

The data on gross migration rates is lacking for Ireland.

Table 2: Cointegration test results

	without trend		with trend	
Belgium	-3,94*	1	-4,45**	1
France	-1,75	1	-2,84	0
Italy	-2,92	4	-3,71	2
Luxembourg	-4,02*	1	-4,81***	5
Netherlands	-3,90*	1	-3,84*	1
Denmark	-3,21	1	-4,33**	4
Ireland	-3,74	1	-3,07	4
UK	-4,12**	1	-3,09	1
Group t-test	-2,31**		-2,27**	

Table reports the country-specific as well as panel cointegration test results. The test statistics of the Engle-Granger cointegration test along with the selected augmentation lag length according to the method suggested in Campbell and Perron (1991) obtained for every country. The 1%, 5%, and 10% critical values when only intercept and no trend were allowed in the model are respectively -4.73, -4.11, and -3.83, and when both intercept and trend were included - -4.65, -4.16, and -3.84, correspondingly. The signs *, **, and *** denote the significant test statistics at the 1%, 5%, and 10% levels, respectively.
The panel cointegration group t-test statistics of Pedroni (1999) has the standard normal asymptotic distribution. The one-sided critical values are -2.32, -1.64, and -1.28 for the respective 1%, 5%, and 10% significant levels.

Table 3: Cointegrating relations

	with trend						without trend					
	$\ln(w_t^*/w_{it})$		$\ln(e_t^*)$		$\ln(e_{it})$		$\ln(w_t^*/w_{it})$		$\ln(e_t^*)$		$\ln(e_{it})$	
Belgium	0.151	*	2.767	***	-1.202	***	-0.255		-6.653	***	0.265	
	(0.08)		(0.45)		(0.27)		(0.28)		(1.27)		(1.00)	
Denmark	-0.400	**	4.914	***	-2.1(64	***	-2.455	***	-2.220		-4.510	***
	(0.20)		(0.79)		0.50)		(0.63)		(2.19)		(1.48)	
France	0.583	***	1.232		1.947	*	0.677	*	3.426		-8.037	***
	(0.12)		(0.82)		(1.13)		(0.35)		(2.32)		(1.87)	
Ireland	1.242	***	2.265		0.423		-0.848		-13.310	*	-5.346	
	(0.16)		(1.80)		(0.80)		(0.66)		(7.64)		(3.56)	
Italy	0.001		2.298	**	4.405	**	0.115	**	2.988	***	-3.236	***
	(0.05)		(1.01)		(1.80)		(0.05)		(1.03)		(1.10)	
Luxembourg	0.279	***	1.775	***	1.886		0.017		-2.217	*	-1.627	
	(0.05)		(0.60)		(2.17)		(0.10)		(1.15)		(4.24)	
Netherlands	-0.151		1.623	***	-0.357		-0.166		1.572	***	-0.409	
	(0.10)		(0.51)		(0.33)		(0.11)		(0.28)		(0.33)	
UK	0.871	***	-0.149		-2.741	***	0.804	***	-15.902	***	1.904	
	(0.10)		(1.71)		(1.05)		(0.21)		(2.37)		(2.17)	

Table reports the slope parameter estimates of equation (2) using the method of canonical cointegrating regressions of Park. Below the estimates, the standard errors are reported in parentheses. The signs ***, **, and * denote the significant test statistics at the 1%, 5%, and 10% levels, respectively.