Migration and the EU Labour Market: Granger Causality Tests on a Panel VAR **

by

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Abstract

This paper investigates the interaction between migration and the European Union (EU) labour market. The study runs Granger causality tests using panel data on 13 of the original EU countries. As predicted by theory, the modelling unveils a negative relationship among unemployment and migration. However, real wages and immigration appear to remain independent. These findings imply that migration flows have had little significant adverse impact on the EU labour market.

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1. Introduction

Migration and associated policies have recently grown in importance as a consequence of Central and Eastern European countries' (CEEC) accession to the European Union (EU). Inter alia, immigration's effect on the employment and real wages of Western Europe's destination countries is an important issue that may involve government restrictions on inflows of potential migrants. Still, there are only a few studies trying to understand this important topic. This paper aims at achieving a better understanding of the effects of migration from the CEECs on the EU's employment and real wages.

The migration literature argues that the higher the perceived probability of finding a job and the higher the real wage in the host country the greater is the rate of migration (Ghatak et al. 1996; Harris and Todaro 1970). In terms of the impact of migration on the labour market in the host country, immigration during recession adds to unemployment either because migrants themselves are unemployed or because native workers are replaced by migrants. However, immigration may also increase aggregate expenditure via migrants' own expenditures and capacity to increase real output by adding to the workforce (Withers and Pope 1985). The standard theory of the labour market predicts that an increase in labour supply exerts a downward pressure on wages. Indeed, the flexibility of the labour market is a major factor in understanding the impact of migration on employment and real wages.

This paper investigates the interaction between migration and unemployment and also between migration and wages by applying Granger's (1969) causality test to panel data on 13 old EU countries for the period 1980-2004. In this regard, it is worth noting that investigation on the relationship between international migration and its impact on the EU's real wages and unemployment is rather limited. Much of the empirical literature concentrates on countries like the USA, Australia and Canada, or on inter-regional migration in transitional economies (Andrienko and Guriev 2004; Ghatak et al. 2007).

The rest of the paper is organised as follows. Section 2 analyses the theoretical relationship between migration and the labour market by focusing on the welfare impact of migration in terms of output and unemployment. Section 3 applies Granger panel causality tests to better understand the problem at hand. Section 4 concludes.

2. A theoretical model of migration

Our theoretical framework to explain migration is based on Harris and Todaro's (1970) model of rural-urban migration, hereafter referred to as H-T. Here, we replace the rural-urban with the East-West; East implies origin country of migrants, whereas West, the destination country of migrants. The future expected income from migration is given by

$$\int_0^\infty \left[PW_W + (1-P)W_b\right]e^{-rt}dt - C = \frac{1}{r}\left[PW_W + (1-P)W_b\right] - C \qquad (1)$$

where C is the direct cost of migration, r is the migrants' discount rate, P is the probability of employment at real wage W_W and W_b is the real income received if unemployed or employed in the informal sector. Would-be migrants compare (1) with the future income from remaining in the East

$$\int_0^{\infty} W_E e^{-rt} dt = \frac{1}{r} W_E \qquad (2)$$

If employment is a certain prospect (i.e. P=1) then migration takes place only if there are gains from moving, i.e., only if

$$\frac{1}{r} W_W - C > \frac{1}{r} W_E \quad or \quad W_W - W_E > rC \qquad (3)$$

Under conditions of uncertainty, the probability of obtaining employment is given by

$$P = \frac{L_W}{N_W} = \frac{L_W}{L_W + M\, N_E} \qquad (4)$$

where L is population employed, N is total population and M is the rate of migrants coming from the East and the subscript $_W$ refers to the West while $_E$ refers to the East. The following equation (5) is derived assuming that equality holds in (3)

$$PW_W + (1-P)W_b - W_E = rC \qquad (5)$$

with P given by (4). Substituting (4) into (5) and solving for M gives the equilibrium migration rate M

$$M = \left[\frac{W_W - W_E - rC}{rC - W_b + W_E} \right] \frac{L_W}{N_W} \qquad (6)$$

We require that $W_b - W_E < rC$ for $M > 0$ which implies that there is an incentive for leaving the East for a West unemployment.

From (6), we get the familiar results

$$\frac{\partial M}{\partial W_W} > 0; \frac{\partial M}{\partial W_E} < 0; \frac{\partial M}{\partial L_W} > 0; \frac{\partial M}{\partial C} < 0 \qquad (7)$$

The conditions in (7) state that marginal increases in the West wage (W_W) or decreases in the East wage (W_E) will increase migration. Paradoxically, any policy to increase employment in the advanced West will raise the migration rate and may increase unemployment in the West. Hence, as predicted in the H-T model, a policy of creating more employment opportunities in the developed counrties may end-up enlarging the migration from the backward countries. Also, any decrease in the cost of migration will increase M. Notably, the H-T model ignores the impact of human capital, availability of public goods like health care, housing stock and road infrastructure on migration decisions (for an alternative model, see Ghatak et al. 2007).

Figure 1 illustrates the migration equilibrium. The real wage/MPL is on the vertical axis and employment on the horizontal axis. Assuming the standard theory of employment determination, initial equilibria occur at G and H, respectively, where the bargained real wages are equal to the MPL in the East (the origin country) and the West (the destination country). Like the standard supply curve of labour, the bargained real wages (BRW) lines slope upwards (Layard et al. 1991). Assume that

OA is the total labour force in East and West prior to migration which is equal. Hence, KA = AB would be equal to migration. Assuming real wage flexibility, the BRW in the West shifts to the right and employment rises by WW'. In the East, BRW shifts to left and employment falls by E'E. The net output gain is given by HJW'W - FGEE' the sign of which cannot be determined a priori. The net change in employment and output will clearly depend on the degree of shifts of the BRW curves in the West. Intuitively, if the labour market and real wages are flexible, the elasticity of real wages with respect to unemployment is high and the unemployment in the West is low, then migration can *increase output and employment.* On the other hand, if the labour market and the real wages in the West are rigid, and the elasticity of the real wages with respect to unemployment is low, then migration can mean loss of output and *increase* in *unemployment* (see Levine 1999). Thus, the overall effect of migration on the job market of the destination country is indeterminate *a priori*. However, the actual measurement of the effect of migration on employment is of considerable research interest. Such an analysis will be undertaken in the next section.

[FIGURE 1]

3. Granger-Causality Tests for the EU using Heterogeneous Panels

To examine the relationship between immigration and the real wages and employment in the EU from 1980 to 2004, we present all the data in logarithmic forms with the following notations: *lmr* = log(migration/population), *lu* = log(unemployment rate) and *lrw* = log(wage rates/prices). The panel data consist of annual observations from 1980 to 2004 (25 time series) for 13 EU member states: Austria, Belgium, Denmark, Finland, France, Germany, Ireland, Italy, Luxembourg, Netherlands, Spain, Sweden and the UK. There is difficulty in obtaining the data for migration, and the choice of country is based on the criteria where more than 18 out of 25 time series are available for the migration data, providing a total of 283 observations. The data for migration is taken from OECD (International Migration Statistics) for the earlier period 1980-84 and Euro-stat for the period 1985-2004. Other data are taken from OECD (Main Economic Indicators) [1].

We employ three types of panel unit root tests, namely of Levin et al. (2002), Im et al. (2003) and Maddala and Wu (1999). Levin et al. (2002) assume that the coefficients on the lagged dependent variables are homogeneous across cross-sections, though incorporates a degree of heterogeneity by allowing for fixed effects and unit specific time trends. The null and alternative are that the whole panel is nonstationary and stationary respectively. In contrast, Im et al. (1997) allows for heterogeneity of the coefficient on the lagged dependent variable with the slope coefficients to vary across cross-sections. The null is that all series are nonstationary while the alternative is that the at least one cross-section is stationary. The Fisher's ADF test proposed by Maddala and Wu (1999) combines the *p*-values from individual unit root tests, and the test statistics are the asymptotic χ^2. See Table 1. The variables in levels are found to be insignificant at the 5% level implying that they are non-stationary. The first

[1] Note that we use gross migration flows instead of net migration flows due to the data availability. Yet, in the study of Jaurez (2000) the variable of gross migration flows is found to perform better than that of net migration in estimation. Also note that although immigrant inflows are small relative to the populations of most countries, recent immigrants are a significant fraction of total low-skilled workers, therefore measuring the effects of immigration at the lower tails of the wage distribution, rather than general wages may generate a different result.

difference of these variables rejects the null of unit root[2]. It follows that the variables are characterised as integrated of order one.

[TABLE 1] and [TABLE 2]

For a panel cointegration test, Pedroni (1999) provides a framework for heterogenous panels by specifying fixed effects, heterogenous slope coefficients and individual specific deterministic trends. The cointegration regression takes the following form without deterministic trends,

$$y_{i,t} = \sum_{i=1}^{N} d_i \alpha_i + \sum_{i=0}^{N} d_i \beta_i lmr_{i,t} + e_{i,t} \tag{8}$$

where y = lu, lrw, N = the number of individual members, T = the number of observation, e = error terms, d_i = 1 for country i, 0 otherwise. The parameter α_i is the country-specific intercept, implying fixed effects and β_i indicates slope coefficients, which vary across individual countries.

Following Pedroni (1999), we have constructed the asymptotic distributions of two panel cointegration statistics based on the residuals of the regression; one is parametric panel *t*-statistics and the other is parametric group *t*-statistics[3]. The asymptotic distributions for the statistics can be expressed in the form of

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[2] Note that Levin et al. (2002) suggests that *lu* in first difference is insignificant at the 5% level, implying the variables is I(2). Yet the superiority of the Im et al. (2003) test and the borderline nature of the Levin et al.'s result seems to justify treating all variables as I(1).

[3] It is assumed that for the parametric panel test, there is a common unit root process so that ρ_{it} in $y_{it} = \alpha + \rho_{it} y_{it-1} + \varepsilon_{it}$ is identical across cross-sections. The group panel test allows for individual unit root processes so that ρ_{it} may vary across cross-sections. The property of the parametric panel test and the group panel test are analogous to that of Levin et al. (2002) and Im et al. (1997) respectively. The specification of panel t-statistics and group t-statistics is found in Table 1 in Pedroni (1999, p.660).

$\frac{k_{N,T} - \mu\sqrt{N}}{\sqrt{v}} \Rightarrow N\,(0,\,1)$ where $k_{N,T}$ is a standardized form with respect to the values of N and T for each statistic and the μ (mean) and v (variance) are functions of the moments of the underlying Brownian motion. The statistics are then compared to the appropriate tails of the normal distribution. The null hypothesis is no cointegration. The empirical results shown in Table 2 fail to reject the null. This implies that there is no long-run equilibrium relationship between unemployment and immigration and also between real wage rates and immigration.

On the basis of the results of no cointegration, one applies the Granger-causality tests by differencing variables. The bivariate panel VAR (PVAR) model takes the following form:

$$\Delta y_t = \sum_{i=1}^{N} d_i \alpha_i + \sum_{j=1}^{\rho} \gamma_j \Delta y_{t-j} + \sum_{j=1}^{\rho} \psi_j \Delta x_{t-j} \qquad (9)$$

where *y*, *x* = *lu* and *lrw* or *lmr* and *lrw* ($y \neq x$), ρ = lag. VAR is normally modelled by maximum likelihood estimation (or GLS) but since the regressors are identical across equations, the estimates are equivalent to those of OLS. Note that in panels, unlike time-series, the use of OLS is only valid or consistent if the time-series dimension (T) is sufficiently large (Binder et al. 2005), and if T is not large enough the application of GMM is suggested. For example, Cao and Sun (2006) and Christiansen and Goudie (2007) assume that around T = 25 is sufficiently large to avoid problems of inconsistency with OLS estimation of the PVAR. On this ground, though T=25 in our case, since data are unbalanced panel data, we apply not only the OLS, but also the GMM. Given a relatively small sample size with annual panel data, which might render sensitivity to the result with the choice of lag, we present three lag lengths of one, two and three, rather than based on selection criteria[4]. The GMM estimation involves instrument variables that are orthogonal to the disturbance terms. Following Binder et al. (2005), instruments used are the levels and lagged two, three and four periods of *lu*, *lrw* and *lmr*.

See Table 3a for the estimates based on OLS. The Hausman test is predominantly rejected in most cases, and so the fixed effects model by specifying α_i is plausible. Although some of the LM second order indicate the presence of serial correlation, all the first order tests suggest the absence of serial correlation for all cases. Where the null of homoskedasticity is rejected, heteroskedasticity is corrected using robust estimation with White's heteroskedastic consistent t-ratios. Table 3b presents the GMM estimates. A similar diagnostic test results seem to apply to the residuals, and the over-identification tests indicate that the null is not rejected in all cases, suggesting that the instruments adopted are valid.

[TABLE 3] and [TABLE 4]

Granger-causality is tested based on the PVAR estimates. See Table 4. In the case of OLS, the $\Delta\, lmr$ Granger-causes $\Delta\, lu$ at all PVAR orders at least at the 10% significance level, whereas for the GMM, it is found to be significant at the 5% level at the first and second orders. As to that $\Delta\, lu$ Granger-causes $\Delta\, lmr$, it is significant at the second and first orders for the OLS and GMM respectively. The results suggest that causality is running in both directions from unemployment to immigration and

[4] If the number of cross sectional observations is small the results may be sensitive to lag length of the PVAR (Christiansen and Goudie, 2007).

from immigration to unemployment. However, wages and immigration seem to be independent of one another.

Table 3a for the VAR model based on the OLS shows that the bilateral relationship between unemployment and migration is negative to each other: the coefficients of Δlmr_{-1} in the Δlu equation and Δlu_{-1} in the Δlm equation are statistically significant with a negative sign. The same results are found for the GMM estimates in Table 3b.

The evidence implies that an increase in the rate of unemployment deters migration being broadly consistent with theoretical and empirical literature. The absence of causality from wages to migration may be explained by the fact that for risk-averse workers, probabilities of employment may be a more important determinant of migration than wage rates (Treyz et al. 1993).

The effect of immigration on unemployment and wages of native workers varies widely from study to study and varies across countries. Our result suggests that the increase of migrants seem to reduce unemployment in the destination countries. The expansionary impact of immigration on employment implies that migration of workers from the new EU does not crowd out national workers, but may have a positive impact on the old EU by alleviating labour market shortages supporting increased employment, which supports the view put forward by the European Commission (2006). As to wages, the result is largely consistent with the recent cross-section studies by Constant (2005) for France and Zorlu and Hartog (2005) for the UK, the Netherlands and Norway, which shows very small migration effects on wages.

4. Conclusion

Employing EU information ranging from 1980 to 2004 and techniques for exploiting panel data this paper concludes that migration from, e.g. CEEC is inversely correlated with unemployment in the destination countries (EU). This conclusion is in line with standard theoretical predictions assuming a flexible labour market. Importantly, the findings imply that the general consensus in the literature that immigrant flows have had little or no substantive adverse impact on the recipient labour market appears to apply in the EU.

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Table 1 Panel Unit Root tests

	Levin, Lin and Chu (2002)		Im, Pesaran and Shin(2003)		ADF - Fisher Chi-square	
	Levels	First differences	Levels	First differences	Levels	First differences
lu	-0.627	-1.595	-0.308	-3.467	29.931	90.271
	[0.265]	[0.055]	[0.379]	[0.000]	[0.270]	[0.000]
lmr	-0.632	-2.543	1.467	-3.097	31.049	121.11
	[0.263]	[0.005]	[0.929]	[0.001]	[0.226]	[0.000]
lrw	0.428	-1.936	0.003	-3.314	25.709	112.80
	[0.665]	[0.026]	[0.501]	[0.001]	[0.479]	[0.000]

Null: unit root. [ ]: prob.

Table 2 Panel Cointegration tests

	Parametric panel *t*-statistics	Parametric group *t*-statistics
Equation (1) (*lu* and *lmr)*	-0.068	-1.275
Equation (2) (*lrw* and *lmr)*	-0.143	-0.843

Null: no cointegration. Critical value -1.64 (5%).

Table 3a Panel VAR Model: OLS

Dep. Var.	$\Delta\ lu$						$\Delta\ lmr$					
	Coef.	t-ratio	Coef.	t-ratio	Coef.	t-ratio	Coef.	t-ratio	Coef.	t-ratio	Coef.	t-ratio
$\Delta\ lu_{-1}$	0.410***	(5.166)	0.409***	(4.099)	0.400***	(4.773)	-0.192***	(2.531)	-0.228***	(3.114)	-0.194***	(2.505)
$\Delta\ lu_{-2}$			-0.001	(0.018)	0.108	(1.102)			-0.061	(0.589)	-0.008	(0.081)
$\Delta\ lu_{-3}$					-0.319***	(4.724)					-0.044	(0.473)
$\Delta\ lmr_{-1}$	-0.077**	(1.914)	-0.095**	(2.244)	-0.104**	(2.305)	-0.003	(0.027)	-0.035	(0.341)	0.042	(0.435)
$\Delta\ lmr_{-2}$			0.071*	(1.706)	0.059	(1.399)			-0.141**	(2.087)	-0.117	(1.558)
$\Delta\ lmr_{-3}$					-0.019	(0.480)					-0.002	(0.028)
	χ^2	[prob.]	χ^2	[prob.]	χ^2	[prob.]	χ^2	[prob.]	χ^2	[prob.]	χ^2	[prob.]
Fixed test	4.387	[0.986]	30.339	[0.004]	10.359	[0.664]	23.746	[0.033]	22.682	[0.045]	16.222	[0.181]
Hausman test	15.731	[0.000]	23.335	[0.000]	30.468	[0.000]	25.718	[0.000]	45.544	[0.000]	43.000	[0.000]
LM serial (1)	0.315	[0.574]	0.537	[0.463]	0.533	[0.465]	0.515	[0.473]	1.839	[0.175]	0.311	[0.576]
LM serial (2)	11.283	[0.003]	12.464	[0.002]	6.067	[0.048]	3.518	[0.172]	6.758	[0.034]	7.579	[0.023]
LM Hetero	1.487	[0.475]	8.470	[0.076]	9.655	[0.139]	12.830	[0.002]	17.980	[0.001]	20.395	[0.002]
Dep. Var.	$\Delta\ lrw$						$\Delta\ lmr$					
	Coef.	t-ratio	Coef.	t-ratio	Coef.	t-ratio	Coef.	t-ratio	Coef.	t-ratio	Coef.	t-ratio
$\Delta\ lrw_{-1}$	0.256***	(3.079)	0.319***	(3.427)	0.319***	(3.082)	-0.006	(0.878)	-0.013*	(1.689)	-0.016*	(1.710)
$\Delta\ lrw_{-2}$			-0.122	(1.244)	-0.120	(1.070)			0.012	(1.083)	0.013	(1.091)
$\Delta\ lrw_{-3}$					-0.080	(1.129)					-0.006	(0.735)
$\Delta\ lmr_{-1}$	-0.124	(0.185)	-0.185	(0.257)	-0.168	(0.270)	0.091	(0.796)	0.069	(0.574)	0.051	(0.453)
$\Delta\ lmr_{-2}$			0.521	(1.013)	0.140	(0.305)			-0.152	(0.974)	-0.185	(1.064)
$\Delta\ lmr_{-3}$					1.337**	(2.099)					0.127	(1.147)
	χ^2	[prob.]	χ^2	[prob.]	χ^2	[prob.]	χ^2	[prob.]	χ^2	[prob.]	χ^2	[prob.]
Fixed test	68.031	[0.000]	55.699	[0.000]	41.333	[0.000]	15.913	[0.253]	14.614	[0.332]	19.714	[0.102]
Hausman test	0.539	[0.763]	29.927	[0.000]	27.588	[0.000]	31.348	[0.000]	45.939	[0.000]	44.973	[0.000]
LM serial (1)	1.145	[0.284]	0.966	[0.325]	1.469	[0.225]	1.318	[0.250]	1.156	[0.282]	1.198	[0.274]
LM serial (2)	3.978	[0.136]	3.767	[0.152]	3.022	[0.220]	5.683	[0.058]	7.398	[0.025]	2.091	[0.352]
LM Hetero	11.316	[0.003]	9.998	[0.041]	5.494	[0.482]	7.778	[0.021]	18.279	[0.001]	24.774	[0.001]

*, ** and ***: Significant at the 10%, 5% and 1% level.

Based on 283 observations of the unbalanced panel data for the period 1980-2004 with 13 EU countries.

Table 3b Panel VAR Model: GMM

Dep. Var.	$\Delta\, lu$ Coef.	t-ratio	Coef.	t-ratio	Coef.	t-ratio	$\Delta\, lmr$ Coef.	t-ratio	Coef.	t-ratio	Coef.	t-ratio
$\Delta\, lu_{-1}$	0.408***	(4.628)	0.418***	(3.924)	0.422***	(4.571)	-0.229***	(2.754)	-0.179**	(2.325)	-0.192**	(2.276)
$\Delta\, lu_{-2}$			-0.003	(0.033)	0.118	(1.160)			-0.012	(0.106)	-0.018	(0.171)
$\Delta\, lu_{-3}$					-0.314***	(4.344)					-0.025	(0.244)
$\Delta\, lmr_{-1}$	-0.092**	(2.128)	-0.093**	(1.972)	-0.097*	(1.808)	-0.056	(0.516)	-0.126	(1.329)	0.097	(0.942)
$\Delta\, lmr_{-2}$			0.084**	(2.053)	0.033	(0.728)			-0.045	(0.632)	-0.030	(0.375)
$\Delta\, lmr_{-3}$					-0.018	(0.430)					-0.088	(0.871)
	χ^2	[prob.]	χ^2	[prob.]	χ^2	[prob.]	χ^2	[prob.]	χ^2	[prob.]	χ^2	[prob.]
Over- id.	1.991	[0.369]	18.666	[0.000]	0.764	[0.682]	1.784	[0.409]	1.008	[0.604]	5.316	[0.070]
LM serial (1)	0.441	[0.506]	0.193	[0.660]	0.143	[0.706]	1.312	[0.252]	1.735	[0.186]	0.311	[0.576]
LM serial (2)	15.429	[0.004]	13.794	[0.001]	5.331	[0.069]	7.427	[0.689]	2.836	[0.242]	7.579	[0.023]
LM Hetero	1.991	[0.369]	7.438	[0.024]	8.758	[0.013]	9.173	[0.010]	7.428	[0.024]	20.395	[0.002]
Dep. Var.	$\Delta\, lrw$ Coef.	t-ratio	Coef.	t-ratio	Coef.	t-ratio	$\Delta\, lmr$ Coef.	t-ratio	Coef.	t-ratio	Coef.	t-ratio
$\Delta\, lrw_{-1}$	0.479***	(6.390)	0.495***	(5.034)	0.454***	(4.392)	-0.256	(0.282)	-0.473	(0.432)	-0.575	(0.506)
$\Delta\, lrw_{-2}$			-0.011	(0.119)	0.004	(0.036)			0.673	(0.581)	1.668	(1.176)
$\Delta\, lrw_{-3}$					0.026	(0.329)					-2.111**	(2.059)
$\Delta\, lmr_{-1}$	-0.004	(0.836)	0.004	(0.671)	0.001	(0.145)	0.099	(0.985)	0.158*	(1.822)	0.142	(1.477)
$\Delta\, lmr_{-2}$			-0.003	(0.693)	-0.003	(0.619)			-0.040	(0.564)	-0.019	(0.259)
$\Delta\, lmr_{-3}$					0.001	(0.087)					0.075	(0.817)
	χ^2	[prob.]	χ^2	[prob.]	χ^2	[prob.]	χ^2	[prob.]	χ^2	[prob.]	χ^2	[prob.]
Over-id.	0.436	[0.804]	1.059	[0.589]	0.262	[0.877]	1.416	[0.493]	4.581	[0.101]	3.919	[0.141]
LM serial (1)	0.117	[0.732]	0.879	[0.348]	1.995	[0.158]	1.189	[0.276]	4.779	[0.029]	2.217	[0.346]
LM serial (2)	0.871	[0.646]	0.735	[0.692]	5.545	[0.062]	0.347	[0.841]	5.113	[0.077]	2.095	[0.351]
LM Hetero	4.237	[0.118]	6.042	[0.049]	6.047	[0.049]	9.397	[0.010]	6.127	[0.047]	13.189	[0.001]

*, ** and ***: Significant at the 10%, 5% and 1% level. Over-id.: Over-identification test

Instrument variables used are *lu*, *lmr*, lu_{-2} and lm_{-2} for the order 1, *lu*, *lmr*, lu_{-2}, lm_{-2}, lu_{-3}, and lm_{-3} for the order 2 and *lu*, *lmr*, lu_{-2}, lm_{-2} lu_{-3}, lm_{-3}, and lu_{-4} lm_{-4} for the order 3.

Based on 283 observations of the unbalanced panel data for the period 1980-2004 with 13 EU countries.

Table 4

a) Panel Granger Causality tests based on OLS

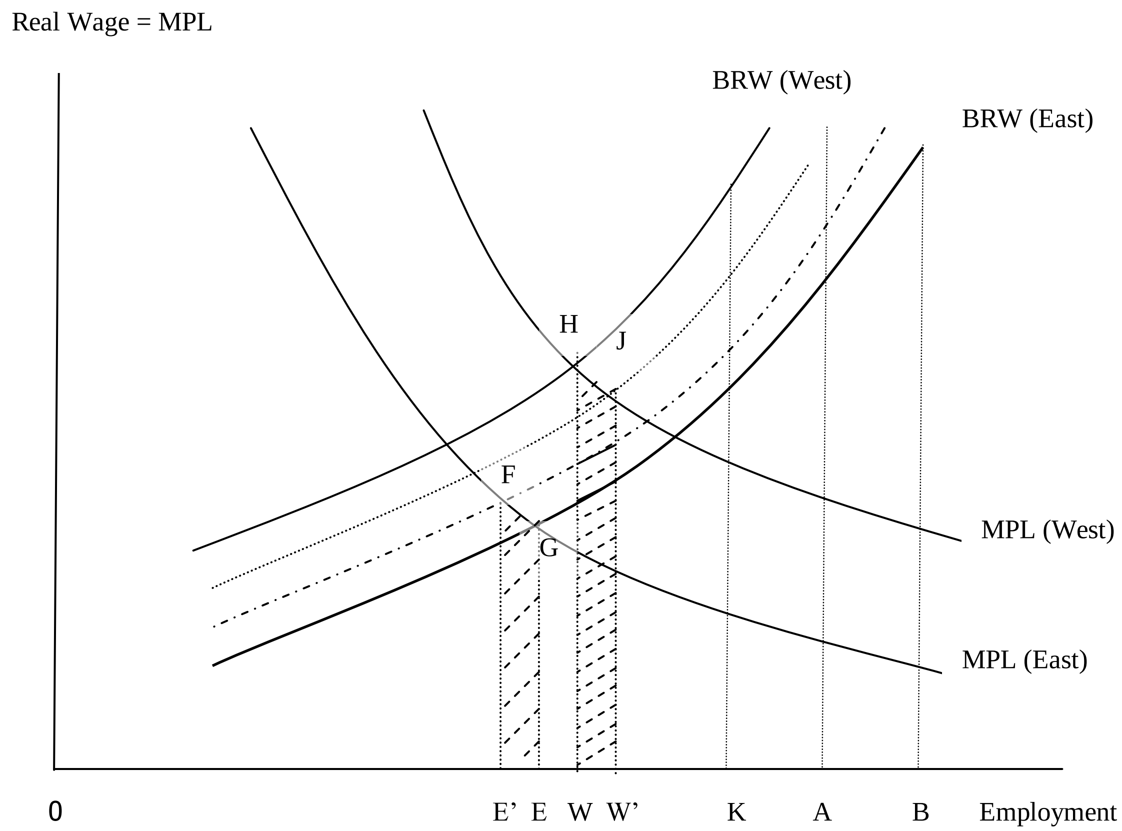
Null hypothesis	PVAR order	GC test χ^2	F test	[Prob]	Results
$\Delta\, lmr$ does not GC $\Delta\, lu$	1	3.662		[0.055]*	Reject null
	2	9.106	4.553	[0.010]***	Reject null
	3	7.808	2.602	[0.050]*	Reject null
$\Delta\, lu$ does not GC $\Delta\, lmr$	1	6.409		[0.011]*	Reject null
	2	10.622	5.311	[0.004]***	Reject null
	3	7.299	2.433	[0.062]	Accept null
$\Delta\, lmr$ does not GC $\Delta\, lrw$	1	0.034		[0.853]	Accept null
	2	1.169	0.585	[0.557]	Accept null
	3	4.720	1.573	[0.193]	Accept null
$\Delta\, lrw$ does not GC $\Delta\, lmr$	1	0.771		[0.379]	Accept null
	2	3.221	1.610	[0.200]	Accept null
	3	3.299	1.099	[0.348]	Accept null

b) Panel Granger Causality tests based on GMM

Null hypothesis	PVAR order	GC test χ^2	F test	[Prob]	Results
$\Delta\, lmr$ does not GC $\Delta\, lu$	1	4.537		[0.033]**	Reject null
	2	8.733	4.367	[0.012]**	Reject null
	3	3.840	1.280	[0.279]	Accept null
$\Delta\, lu$ does not GC $\Delta\, lmr$	1	7.588		[0.005]***	Reject null
	2	5.848	2.924	[0.053]*	Reject null
	3	5.617	1.873	[0.132]	Accept null
$\Delta\, lmr$ does not GC $\Delta\, lrw$	1	0.698		[0.403]	Accept null
	2	0.695	0.348	[0.706]	Accept null
	3	0.399	0.133	[0.940]	Accept null
$\Delta\, lrw$ does not GC $\Delta\, lmr$	1	0.079		[0.778]	Accept null
	2	0.359	0.179	[0.836]	Accept null
	3	4.258	1.419	[0.234]	Accept null

[ ]: prob.   *, ** and ***: Significant at the 10%, 5% and 1% level.

Figure1. Migration, output and employment



BRW = Bargained real wage
OA = Total labour force in East and West prior to migration
KA = AB = Migration
BRW (West) shifts to right: Employment rises by WW'
BRW (EAST) shifts to left: Employment falls by E'E
Net output gain = HJW'W – FGEE' of indeterminate sign
Net employment gain: indeterminate sign