COINTEGRATION MODELING OF INTERRELATED FACTOR DEMANDS:
With an application to labor-import substitution in the European Union

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Abstract

In this paper we present techniques for cointegration modeling of interrelated factor demands. These techniques respect the non-stationary character of the price and quantity data, and permit specification of general, dynamic factor demand models based on error correction forms derived from cointegration. Therefore, we do not have to assume *ad hoc* dynamic forms. Moreover, we ensure that estimated relations are structural, and not spurious. Cointegrating vectors are estimated subject to all standard economic theory restrictions by using a procedure, which we call dynamic SUR. We show how consistent error correction models can be specified and estimated. In addition, we test the neoclassical restrictions both in the short- and the long run. The new methods are used to shed light on the major problem of the European Union (unemployment) and its relationship with imports. The empirical analysis is conducted for five countries of the European Union with an emphasis to the south: The UK, France, Greece, Italy, and Spain.

Key words: Cointegration, error correction, dynamic SUR, interrelated factor demands, dynamic adjustment.

JEL codes: C22, F10, O57.

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1. Introduction

Several economic time series have stochastic trends, and many of them could be cointegrated. Therefore, we expect that certain economic relationships are structural, and not spurious. Therefore, cointegration modeling would be the proper way of approaching the problem of dynamic interrelated factor demands, it has not been fully studied to permit routine use in applied studies. One reason, is undoubtedly, that it is not obvious how the usual share equation systems translate into cointegrating relations that could be estimated by Johansen [1991] style maximum likelihood procedures. Although it is clear that log-linear (*i.e.* translog) share equations form structural, cointegrating relations, it is not clear how they can be estimated *jointly* subject to *all the restrictions* imposed by the theory. It is also unclear how error correction forms, consistent with the long-run cointegrating restrictions can be formulated and estimated.

Empirical studies on interrelated equation systems for factor or consumer demand such as system of share equations, are mainly based on static specifications or specifications that are dynamic but in an *ad hoc* way. In static models, the implicit assumption is that factor shares adjust instantaneously to desired levels following exogenous price and demand shifts. Thus, these models do not provide the adjustment path. Information about the adjustment process, however, is often important in addressing policy issues. *Ad hoc* specifications have the drawback that they are not consistent with a consistent theory of adjustment and thus, they may be difficult to accept.

Early contributions to the theory of dynamic specification of interrelated factors of production include Anderson and Blundell [1982] and Nissim [1984]. Anderson and Blundell [1982] model the dynamic structure by transforming a general autoregressive distributed lag system of share equations to an error correction specification. In an analysis of labor demand, Nissim [1984] used a similar model but for factor demands. Friesen [1992] has applied the Anderson and Blundell [1982] approach to estimate factor demand functions for US manufacturing and he found that it is dominates several more parsimonious models. However, he rightly concludes that *the issue of dynamic specification remains unresolved.*

Jones [1995] extended the dynamic version of share systems by proposing a dynamic logit model to investigate inter-fuel substitution in US industrial energy demand. For other applications, see Ryan and Wales [2000] and Tombazos [1998]. The latter study concerns estimation of a model for the US containing capital, labor, and imports, so it is related in spirit to this study. However, the method of estimation is static SUR, which ignores the dynamics and, of course, the integration-cointegration properties of the data. Allen and Urga [1999] proposed an *ad hoc* dynamic specification of a cost function - factor demand system. The implication is that previous research proposed general dynamic models for factor demands but did not provide a consistent relationship between the short run and long run factor demands. In other words, it has not been possible so far to explore the implications of cointegration for factor demand systems, and the implications of cointegration for the short-run dynamics of factor demands.

In this paper, a practical estimation procedure is proposed based on the dynamic OLS approach of Stock and Watson [1993] extended to allow for factor demand systems of equations. The approach takes into account the cointegration properties of the underlying time series as well as the dynamic properties of residuals, and can be easily implemented using standard econometric packages. The new technique also produces an error correction form for all factors of production that is consistent with the cointegrating vector, and the adjustment to long-run equilibrium without the need to resort to *ad hoc* dynamic specifications.

The new techniques are applied to an investigation of factor demands for capital, labor, and imports in five countries of the European Union to determine whether or not imports contribute to unemployment. The issue of unemployment is currently the major problem of the European Union. Indeed, lower productivity growth compared to the US, and several institutional problems are thought to be responsible for the continual depreciation of the euro relative to the US dollar: This masks the structural weaknesses of the European economy, the most important of which are undoubtedly productivity growth gaps, and unemployment. Moreover, it is well known that the European South lags behind in terms of productivity and competitiveness relative to the rest of the European Union. Therefore, a productivity shock that increases domestic prices relative to the competitors will trigger an increase in imports of goods and services– not to mention the dependence on oil imports. The employment effects depend on the composition of imports (the extent to which they are imports of final or intermediate goods). Consequently, there is an empirical issue concerning the substitutability or complementarity between imports and employment in the European Union, and particularly the southern component. Moreover, according to the Stolper-Samuelson theorem in a two-good (say labor and capital), two-factor, competitive world with no joint production, the imposition of a tariff redistributes real income from one factor to the other. Apparently, the direction of the re-distribution effect depends crucially on the substitution elasticities of labor with imports and capital with imports, since a tariff affects the demand for imports. These elasticities must be estimated in order to make statements about the re-distribution of real income that takes place within the European Union as the result of free trade.

The rest of the paper is organized as follows. The theoretical model is presented in section 2. The estimation technique is described in section 3. The empirical application as well as the discussion of results, and the policy implications are in section 4. The final section concludes the paper.

2. Model specification

The technology of production can be described by a transformation function $f(Y;X)=0$ where $Y=[Y_1,Y_2....Y_m]$ is the vector of outputs and $X=[X_1,X_2....X_n]$ is the vector of inputs. We would like a functional form that allows for an arbitrary set of partial elasticities of substitution and transformation between pairs of inputs and outputs. These functions can offer substantial gains over the traditional Cobb-Douglas or CES functions. Christensen, Jorgenson and Lau [1973] have proposed a functional form for the transformation function known as translog which is second-order approximation to an arbitrary cost or production function. In this paper we assume that the production technology can be described by a translog function form.

If P_K is the price of capital, P_L is the price of labor, P_M is the price of imports, T is a time trend, and Y is output, the translog minimum cost function adopted in this study is given by:

$$\ln C_t = \alpha_o + \sum_i \beta_i \ln P_{it} + 0.5\sum_i \sum_j \beta_{ij} \ln P_{it} \ln P_{jt} + \sum_j \rho_{iY} \ln P_{it} \ln Y_t + \sum_j \delta_{it} \ln P_{it} t + \\ + \sum_j \delta_{Yt} \ln Y_t t + \alpha_Y \ln Y_t + 0.5 \ln Y_t^2 + \alpha_t t + 0.5\alpha_{TT} t^2 \qquad [1]$$

where $i, j = K, L, M$. Differentiating [1] with respect to factor prices and applying Shephard's lemma the cost share equations are obtained:

$$S_{it} = \frac{\partial \ln C_t}{\partial \ln P_{it}} = \beta_i + \sum_j \beta_{ij} \ln P_{jt} + \rho_{iY} \ln Y_t + \delta_{it} t \qquad [2]$$

To satisfy the neoclassical properties, the following linear restrictions must be satisfied:

$$\sum_i \beta_i = 1,\ \sum_i \beta_{ij} = \sum_j \beta_{ji} = 0,\ \ \sum_i \rho_{iY} = 0,\ \sum_i \delta_{it} = 0 \qquad [3]$$

Restrictions [3] imply that the cost function is linearly homogeneous in factor prices. In the translog cost function, long run (LR) own-price elasticities and the Allen Uwaza partial elasticities are defined as follows:

$$\varepsilon_{ii}^{LR} = \frac{\beta_{ij} + S_i^2 - S_i}{S_i^2},\ \sigma_{ij}^{LR} = \frac{CC_{ij}}{C_i C_j} \text{ where } C_i = \frac{\partial TC}{\partial P_i},\ C_{ij} = \frac{\partial^2 (TC)}{\partial P_i \partial P_j} \qquad [4]$$

More specifically, $\sigma_{ij}^{LR} = 1 + (\frac{\beta_{ij}}{S_i S_j})$. [5]

3. The DSUR method and error correction forms

Factor demand systems of the translog variety can be expressed as linear equation systems subject to cross-equation restrictions. We consider the following system of equations

$$\mathbf{y}_{1t} = \boldsymbol{\mu}_1 + \boldsymbol{\Theta}\mathbf{y}_{2t} + \mathbf{u}_{1t} \qquad [6]$$

$$\mathbf{y}_{2t} = \mathbf{y}_{2,t-1} + \boldsymbol{\mu}_2 + \mathbf{u}_{2t} \qquad [7]$$

$$\begin{bmatrix} \mathbf{u}_{1t} \\ \mathbf{u}_{2t} \end{bmatrix} = \mathbf{A}(L)v_t = \sum_{s \geq 0} \mathbf{A}_s \mathbf{v}_{t-s} \qquad [8]$$

$$E(\mathbf{v}_t) = \mathbf{0},\ E(\mathbf{v}_t \mathbf{v}_t') = \boldsymbol{\Omega},\ E(\mathbf{v}_t \mathbf{v}_s') = \mathbf{0}\ (t \neq s) \qquad [9]$$

where $\mathbf{y}_{1t}$ and $\mathbf{y}_{2t}$ are $n \times 1$ and $m \times 1$ vectors of variables. In production studies, $\mathbf{y}_{1t}$ typically represents the vector of shares, and $\mathbf{y}_{2t}$ the vector of log prices and log output. In demand studies, $\mathbf{y}_{1t}$ would be the vector of shares of goods, and $\mathbf{y}_{2t}$ the vector of log prices and log income.

As long as the error terms $\mathbf{u}_{1t}$ and $\mathbf{u}_{2t}$ are uncorrelated, it is well known that despite integration and cointegration, the correct approach is to estimate the first equation by OLS because the distributions of parameter estimators and Wald statistics, are standard. This approach has been implemented by Attfield [1997] in the context of consumer demand systems of the AIDS variety.

The critical condition of this approach is that $\mathbf{A}(L)$ can be factored as $\mathbf{A}(L)=\begin{bmatrix} a_{11}(L) & \mathbf{0}' \\ \mathbf{0} & \mathbf{A}_{22}(L) \end{bmatrix}$. When this assumption cannot be maintained, and there is a single cointegrating vector ($\mathbf{y}_{1t}$ is a scalar), Phillips and Loretan [1991], and Stock and Watson [1993] have suggested the method of dynamic OLS which adds lags and leads of Δy_{2t} to the first equation. The spirit of this approach can be followed here. In other words, we consider the system

$$\mathbf{y}_{1t} = \boldsymbol{\mu}_1 + \boldsymbol{\Theta}\mathbf{y}_{2t} + \sum_{s=-p}^{p} \boldsymbol{\Gamma}'_s \Delta\mathbf{y}_{2,t-s} + \tilde{\mathbf{u}}_{1t} \qquad [10]$$

under the assumption that $\tilde{\mathbf{u}}_{1t}$ and $\tilde{\mathbf{u}}_{2t}$ are uncorrelated at all lags and leads. Since there are cross-equation restrictions among the elements of $\boldsymbol{\Theta}$, the appropriate method of estimation is SUR subject to the appropriate restrictions. One problem is autocorrelation of errors, which could be handled in a number of ways: Stock and Watson [1993] suggest the correction of standard errors and t-statistics by factors that depend on the estimated autocorrelations from the cointegrating regressions. Phillips and Loretan [1991, p. 424] suggest the inclusion of lags of $\mathbf{y}_{1t} - \boldsymbol{\Theta}\mathbf{y}_{2t}$ in the system equations to be estimated. This approach results in the following system of equations:

$$\mathbf{y}_{1t} = \boldsymbol{\mu}_1 + \boldsymbol{\Theta}\mathbf{y}_{2t} + \sum_{s=-p}^{p} \boldsymbol{\Gamma}'_s \Delta\mathbf{y}_{2,t-s} + \sum_{s=1}^{p} \boldsymbol{\Psi}_s(\mathbf{y}_{1t} - \boldsymbol{\Theta}\mathbf{y}_{2t}) + \tilde{\mathbf{u}}_{1t} \qquad [11]$$

where $\boldsymbol{\Psi}_s$ is an $n \times n$ matrix of parameters. The approach proposed here deals effectively with four problems: Cointegration among the variables, the necessity to impose cross-equation restrictions derived from economic theory, simultaneity, and autocorrelation of error terms. This method will be called dynamic SUR, or DSUR for brevity.

Standard econometric theory implies that the error correction model (ECM) for the cost share equations in [2] is given by:

$$\Delta S_i = \sum_{m=1}^{q} \alpha_{im} \Delta S_{i,t-m} + \sum_{m=0}^{q} \gamma_{ijm} \Delta \ln P_{j,t-m} + \sum_{m=0}^{q} \theta_{im} \Delta \ln Y_{t-m} + \phi_i EC_{i,t-1} + e_{it} \qquad [12]$$

$$i = K, L, M \quad t = q+1, \ldots, T$$

where Δ is the first difference operator, $EC_{i,t}$ denotes the error correction term of the i*th* share (which represents the lagged value of the residual from the cointegrating regression), q is the number of lags, and e_{it} is statistical noise. Parameters ϕ_i are the speeds of adjustment to long-run equilibrium.

An important issue is elementary consistency of the error correction form in [12]. Specifically, we must have $\sum_i \Delta S_{it} = 0$ for all t, so the following restrictions are introduced:

$$\alpha_{im} = \alpha_{Km} \text{ for all } i, m, \; \sum_i \gamma_{ijm} = 0 \text{ for all } j, m, \; \sum_i \theta_{im} = 0 \text{ for all } m \qquad [13]$$

$$\phi_i = \phi_K \text{ for all } i, \; \sum_i EC_{it} = 0 \text{ for all } t \qquad [14]$$

where $i, j = K, L, M$, $m = 0,1,\ldots,q$, and $t = 1,\ldots,T$.

In particular, these restrictions require that the speeds of adjustment, ϕ_i, must be the same for all factors of production, and that the average (across factors) of error correction is zero. Otherwise, time-differenced shares could not sum to zero for each time period. Restrictions [13] and [14] must be imposed in the short run to ensure elementary consistency of the error correction representations in [12].

4. Empirical results

4.1 Data, stationarity and cointegration tests

We use annual data for the UK, France, Greece, Italy, and Spain for 1960-1998. Data for capital, labor, imports, output, and factor prices were obtained from the AMECO database (for the nature and structure of the variables see the Appendix). To test for stationarity, we have used the augmented Dickey-Fuller (ADF) and the Phillips-Perron (PP) tests. The results are presented in Table 1, and indicate clearly that in all five countries, variables P_i, Y and S_i, i=K,L,M contain a unit root. We also provide augmented ADF and PP tests for all series in first differences. The result is that, all series are stationary in their first differences except of labor prices in France and Italy, and output series in France and Spain, where the ADF statistic indicates the presence of a unit root in first differences. According to the Phillips-Perron statistics, however, this is decisively not the case. Therefore, we can proceed on the working hypothesis that all series are I(1).

[Insert Table 1]

To test for *cointegration* the maximum likelihood methodology developed by Johansen [1988] is used. Cointegration results are shown in Table 2, which report Johansen maximum eigenvalue test statistics. The critical values are adjusted for the number of variables, number of lags and number of observations following the methodology presented in Cheung and Lai [1993]. The results show that labor and capital equations for each country have exactly one cointegrating vector at the conventional 5% level. Therefore, our findings confirm the presence of a long run equilibrium relationship between p_i, y and S_i, i=K,L,M, for all countries. This is important, because if cointegration were rejected, share equations would not constitute structural but spurious relationships.

[Insert Table 2]

4.2. Elasticities of substitution

In Table 3 we present DSUR estimates of the share equations [6] and [7] with linear homogeneity and symmetry imposed in the long run. The regression results indicate that almost all long run coefficients are statistically different from zero at conventional levels of statistical significance. Further, Table 3 also reports various diagnostic statistics for the specification of our estimated model. Our results indicate that the DSUR model is well specified and free from specification errors.

Since Greece and Spain joined the EU in different periods (Greece joined the EU in 1981 and Spain in 1986) there is the possibility of a structural shift in the relationship between factor shares, factor prices and output. In other words an issue of structural stability arises. If that is correct then the model parameters may have changed thus leading to inconsistent elasticity estimates. We examine the structural stability of the equations by testing for constancy of the intercepts. Tests involving the slopes would be unreliable due to the small number of degrees of freedom. The tests do not provide evidence against the stability of the model (the computed χ^2 for the equation of capital in Greece and Spain is 0.56 and 1.63 respectively with *p*-values 0.45 and 0.50 while for the equation of labor the χ^2 values are 0.29 and 2.12 respectively with p-values 0.29 and 0.34).

[Insert Table 3]

Next we substituted the estimated coefficients of share equations from Table 3 in the [4] and [5] to derive estimates of *long run elasticities* for each of the production factors, capital (K), labor (L) and imports (M). The results are shown in Table 4. Overall, our results suggest that all own price elasticities are negative as expected, and all of them are statistically significant at the usual levels of significance. Partial elasticities of substitution between capital and labor are positive indicating that these two factors are substitutes. They range from 0.75 and 0.77 for Greece and France to 1.01 for the UK. Their standard errors are small, indicating that the DSUR method provides accurate point estimates of these elasticities.

Also substitutes are capital and imports. These elasticities are reasonably close to unity, ranging from 0.91 for Greece to 1.11 for Spain, while for the UK the elasticity is 1.88 and highly statistically significant. The most interesting elements in Table 4 from our point of view, are elasticities between labor and imports. With the exception of the UK, these elasticities indicate that labor and imports are complements. They are very high for France and Italy (-0.83 and –0.91 respectively), followed by Greece (-0.59) and Spain (-0.32). For the UK this elasticity is close to unity (1.01) and highly significant. For Greece and Spain we cannot in fact reject the hypothesis that labor and imports are unrelated. So only for France and Italy we can document that labor and imports are complements, while in the UK they are substitutes.

Aw and Roberts [1985] also found a complementarity relationship between labor and imports in the US. This result (a reminder of Leontief's paradox) *contradicts the predictions of traditional trade theory that assumes that all traded goods are final*. In this case the main bulk of imports consists of intermediate goods which may stimulate the demand for labor, by a grater amount than they decrease it, via substitution of import-competing commodities, resulting in a net positive effect.

[Insert Table 4]

The important policy implication is that in the South (Greece and Spain) an increase in imports does not have an employment effect. On the contrary, in France and Italy we expect a positive effect and a negative effect for the UK. These results support the view that for the non-industrial Southern European countries, namely Greece and Spain more openness and trade will not affect positively employment. Therefore, it does appear to be true that a Europe of "two speeds" will emerge in the long run because of the very different characteristics of the economies. Certain members of the European Union like Italy and France are, of course, affected positively by trade in employment terms, while Spain and Greece are not affected at all.

Therefore, more trade will contribute positively, to some extent, to the unemployment problems of France and Italy. The reason why this does not happen in the UK, is that the structure of imports is very different in the industrial economy of the UK, and the economies of Italy and France. For example, if the bulk of imports consists of intermediate goods requiring further processing it is not surprising that we find complementarity in Italy and France but not in the UK. Therefore, at the fundamental level industrial structure has a lot to do with the employment-imports relation.

Finally, we compare our results with those reported by other researches, see Table 5. Relative to previous studies we are in general agreement regarding the sign of partial elasticities of substitution. Some relevant papers are Apostolakis [1981, 1988] and Truett and Truett [2000]. For σ_{KL}, results for Italy, Spain and UK are about the same in terms of magnitude while some differences arise for σ_{KM} and σ_{LM}. For the UK all elasticities are about the same compared to Apostolakis [1988]. Regarding σ_{LM} for Italy and Spain we obtain complementarity contrary to Apostolakis [1981] and Truett and Truett [2000] save, of course, for the UK where we agree with Apostolakis [1981] that σ_{LM} indicates substitutability. Overall, although we are in general agreement with previous research about the sign and magnitude of σ_{KM} we cannot say the same regarding the important parameter of this paper, namely σ_{LM}. Naturally, this can be attributed to the improved estimation technique that respects the non-stationarity character of the data, and the use of DSUR techniques. This is critical because the sign of σ_{LM} is critical for important practical issues like the effect of trade on unemployment.

[Insert Table 5]

4.3. Error correction models

Although the above discussion is valid in the long run, one may argue that in the short run the situation may be different. To derive *short run elasticities* we estimate the error correction models [12] via GMM estimation techniques. GMM estimation is necessary because we include the current value of own price log changes in system [12]. We compute standard errors robust for autocorrelation and heteroscedasticity[1]. In this case OLS estimation produces inconsistent estimators. Further, in the aggregate level, the standard assumption that prices and output are exogenous may not hold, so a method like GMM must be used. To this end, we perform standard Johansen exogeneity tests. The results are reported in Table 6 and indicate that prices can not be considered as exogenous variable for Spain and France and output series cannot be considered exogenous for the UK. Therefore GMM will be an appropriate estimation technique for our VEC model.

To implement GMM we have used as instruments the exogenous variables of the model, namely prices and outputs lagged L periods. We started assuming $L = 2$. The choice of L turned out to be somewhat important in our context, because for France and Italy with $L = 2$ the estimate of ϕ was positive contradicting that we have short-run adjustment to the long run shares. Increasing $L = 4$ gave acceptable results as reported below. For other countries, increasing L from 2 to 4 did not produce significantly different results.

Before we proceed to the estimation of VEC models [12] we need to test the existence of the neoclassical restrictions in the short run to ensure consistency of the error correction representations in [12]. In Table 7 we report tests of the neoclassical restrictions in the short run. These restrictions cannot be rejected. In particular, we find that the restriction of common adjustment speeds for all factors of production cannot be rejected, and this is true for all countries. Indeed, the significance levels for these restrictions range from 0.18 for Greece to 0.91 for France.

[Insert Tables 6 and 7]

GMM estimates as well as diagnostic statistics for the vector error correction models are reported in Table 8. According to these results the hypothesis of normal standard errors cannot be rejected while RESET tests rejects that there is functional form misspecification. Again, an issue of structural stability arises for Greece and Spain. Like in the case of DSUR, the results are in favor of the stability of the model. The computed χ^2 for the equation of capital in Greece and Spain is 1.97 and 2.05 respectively with p-values 0.16 and 0.15 while for the equation of labor the χ^2 values are 0.001 and 1.08 respectively with p-values 0.97 and 0.30.

The most important information in Table 8, concerns estimates of adjustment speeds to a long run shock namely the ϕ_i parameters. These parameters are negative, indicating that upward deviations from the long run equilibrium are followed by downward corrections in the short run to compensate for these deviations. As mentioned previously, the choice of the number of lags for the instruments was somewhat important in obtaining valid short-run adjustment equations for France and Italy but for other countries that was not an issue.

[Insert Table 8]

In Table 9 reported are the short run elasticities derived from the VEC models [12]. The signs of elasticities between labor and imports remain the same such in the long run as well in the short run period. The only major difference that emerges between the two periods is that only Italy now has statistically significant cross price elasticity. In the short run, however, the elasticity is much lower (-0.37 compared to –0.91 in the long run). These findings suggest that the employment effect of imports is negligible in the European South with the exception of Italy, but very significant and positive in the UK, both in the short- and the long run.

Since $\sigma_{LM} < 0$, both in the short-run and the long-run for all countries we consider, save for the UK, higher tariffs must raise the remuneration to factors of production that are complementary to imports. If tariffs decline, as should be the case with free trade within the European Union, since $\sigma_{KM} > 0$ income is re-distributed from workers to capitalists. The re-distribution effect is a well known implication of the Stolper-Samuelson theorem according to which in a two-good, two-factor competitive world with non-joint production, after the imposition of a tariff, the real income of one factor necessarily increases, and the real income of the other factor necessarily decreases, see also Apostolakis [1990].

5. Conclusions

In this paper we have described how interrelated factor demand models can be estimated by the dynamic SUR method in a way that respects the integration and cointegration properties of the underlying time series, and produces error correction forms for all factors of productions. The method can be used to deliver estimates of long run as well as short run price elasticities and can be implemented in standard software. The contribution of the paper is that it delivers the implications of cointegration for factor demand systems, and the implications of cointegration for the short-run dynamics of factor demands. This is important because so far dynamic models for factor demands have been on an *ad hoc* basis. From the econometric point of view, the new approach deals effectively with four problems: Cointegration among the variables, the necessity to impose cross-equation restrictions derived from economic theory, simultaneity of factor demands, and autocorrelation of error terms

We have applied the new methods to an examination of import-labor substitution in five countries of the European Union, with an emphasis to the South, in order to examine whether imports contribute to unemployment, the major problem of the European economy today. Our framework was a system of interrelated factor demands consisting of labor, capital, and imports. The empirical results suggest that only for France and Italy we can document that labor and imports are complements, while in the UK they are substitutes. The important policy implication is that for the South (Greece and Spain) an increase in imports does not have a statistically significant employment effect. In France and Italy we do have a positive employment effect, and a negative employment effect for the UK. Therefore, these results support the point of view that for the non-industrial Southern European countries, more openness and trade will not affect positively employment. The positive effects of imports are likely to be contained within the major industrial countries of the European Union with the exception of the UK. Taking account of short run information from error correction models, it turns out that only Italy has statistically significant cross price elasticity. Thus the employment effect of imports is negligible in the European South with the exception of Italy, but very significant and negative in the UK, both in the short- and the long run.

Table 1. Dickey-Fuller (ADF) and Phillips-Perron (PP) tests for unit roots

Country	Tests	$\ln P_K$	$\ln P_L$	$\ln P_M$	$\ln Y$	S_K	S_L	S_M
Greece	*ADF (L)*	**-0.89**	**-1.91**	**-1.31**	**-2.46**	**-2.11**	**-2.85**	**-2.00**
	PP(L)	**-1.32**	**-1.93**	**-1.04**	**-2.29**	**-0.98**	**-2.69**	**-1.46**
	ADF (FD)	-3.64[a]	-3.65[a]	-3.44[b]	-3.78[b]	-4.63[a]	-4.46[a]	-3.42[b]
	PP (FD)	-7.83[a]	-4.29[a]	-4.10[a]	-5.53[a]	-8.07[a]	-6.01[a]	-6.34[a]
France	*ADF (L)*	**-1.33**	**-2.15**	**-1.20**	**-2.67**	**-1.34**	**-1.78**	**-1.66**
	PP(L)	**-1.30**	**-2.33**	**-1.02**	**-2.53**	**-1.45**	**-2.32**	**-1.87**
	ADF (FD)	-3.40[c]	**-2.40**	-3.21[c]	**-2.98**	-3.60[b]	-4.99[a]	-5.28[a]
	PP (FD)	-3.92[b]	-2.69[c]	3.81[b]	-4.10[b]	-5.36[a]	-7.26[a]	-8.51[a]
Italy	*ADF (L)*	**-1.13**	**-2.63**	**-1.26**	**-1.56**	**-2.27**	**-2.99**	**-1.71**
	PP(L)	**-0.85**	**-2.81**	**-1.19**	**-2.19**	**-1.87**	**-2.94**	**-2.18**
	ADF (FD)	-4.23[a]	**-2.78**	-3.41[c]	-4.80[a]	-4.02[b]	-4.41[a]	-5.50[a]
	PP (FD)	-5.67[a]	-3.48[b]	-3.69[b]	-5.46[a]	-5.35[a]	-7.58[a]	-7.71[a]
Spain	*ADF (L)*	**-1.62**	**-0.09**	**-1.18**	**-2.29**	**-2.81**	**-1.19**	**-2.35**
	PP(L)	**-1.70**	**-0.17**	**-1.06**	**-2.36**	**-2.24**	**-1.35**	**-2.20**
	ADF (FD)	-3.48[b]	-3.54[b]	-3.74[b]	**-2.78**	-3.21[c]	-4.13[b]	-4.31[a]
	PP (FD)	-4.78[a]	-4.55[a]	-4.00[b]	-3.93[b]	-4.26[a]	-4.72[a]	-5.48[a]
UK	*ADF (L)*	**-2.60**	**-1.92**	**-1.34**	**-2.46**	**-2.73**	**-2.13**	**-2.83**
	PP(L)	**-1.31**	**-1.69**	**-1.32**	**-2.49**	**-1.80**	**-2.29**	**-3.06**
	ADF (FD)	-5.65[a]	-4.42[a]	-3.92[b]	-4.64[a]	-5.54[a]	-5.58[a]	-5.49[a]
	PP (FD)	-4.25[a]	-4.14[b]	-4.49[a]	-4.36[a]	-4.29[a]	-7.26[a]	-6.95[a]

Notes*: ADF (L) and PP (L) denote the augmented Dickey-Fuller and Phillips-Perron t-tests for a unit root in the levels model (constant/trend) respectively. ADF (FD) and PP (FD) apply to the first differences model. Number of lags was selected using the Schwarz criterion. Boldface values denote sampling evidence in favor of unit roots. (a), (b) and (c) signify rejection of the unit root hypothesis at the 1%, 5% and 10% level respectively.*

Table 2. Johansen cointegration tests. Max eigenvalue statistics for H_0: rank=r

Labor equation

Country	*R=0*	*r≤1*	*r≤2*	*r≤3*
Greece	84.34*	37.55	17.23	7.35
	[70.05]	[48.45]	[30.45]	[16.26]
	[75.64]	**[52.33]**	**[32.86]**	**[17.56]**
France	111.96*	59.13	35.35	5.77
	[87.31]	[62.99]	[42.44]	[25.32]
	[93.42]	**[67.39]**	**[45.41]**	**[27.09]**
Italy	69.49**	40.73	21.88	8.75
	[62.99]	[42.44]	[25.32]	[12.25]
	[67.44]	**[45.41]**	**[27.09]**	**[13.72]**
Spain	89.19*	41.09	16.28	5.82
	[70.05]	[48.45]	[30.45]	[16.26]
	[75.64	**[52.33]**	**[32.86]**	**[17.56]**
UK	77.86*	25.02	11.10	3.02
	[55.46]	[35.65]	[20.04]	[6.65]
	[69.32]	**[44.56]**	**[25.05]**	**[8.31]**

Capital equation

Country	*r=0*	*r≤1*	*r≤2*	*r≤3*
Greece	85.24*	42.17	22.89	8.57
	[70.05]	[48.45]	[30.45]	[16.26]
	[75.64]	**[52.33]**	**[32.86]**	**[17.56]**
France	76.72**	38.69	22.14	9.45
	[62.99]	[42.44]	[25.32]	[12.25]
	[67.39]	**[45.41]**	**[27.09]**	**[13.72]**
Italy	75.35**	41.33	22.89	8.93
	[62.99]	[42.44]	[25.32]	[12.25]
	[67.39]	**[55.41]**	**[27.09]**	**[13.72]**
Spain	85.35*	33.63	16.89	7.15
	[70.05]	[48.45]	[30.45]	[16.26]
	[75.64]	**[52.33]**	**[32.86]**	**[17.56]**
UK	72.15*	27.42	9.86	2.68
	[55.46]	[35.65]	[20.04]	[6.65]
	[69.32]	**[44.56]**	**25.05]**	**[8.31]**

Notes*: r denotes the number of cointegrating vectors. The optimal lag structure for the VAR was selected by minimizing the Schwarz criterion. Figures in brackets are critical values.* and ** indicate rejection of the non-cointegration hypothesis at the 1% and 5% level of statistical significance respectively. Bold face values are adjusted critical values for the number of variables, number of lags and number of observations, see Cheung and Lai[1993].*

Table 3. DSUR estimation of share equations
(homogeneity and symmetry restrictions are imposed)

Param.	*Greece*	*France*	*Italy*	*Spain*	*UK*
β_K	0.648* [0.008]	0.541* [0.06]	0.542* [0.008]	0.563* [0.008]	0.471* [0.002]
β_{KK}	0.046* [0.006]	0.035* [0.003]	0.009* [0.004]	0.006 [0.003]	-0.998* [0.004]
β_{KL}	-0.036* [0.006]	-0.041* [0.005]	-0.024* [0.006]	-0.013* [0.004]	-0.002 [0.004]
ρ_{KY}	-0.035** [0.016]	-0.179* [0.011]	-0.228* [0.028]	-0.164* [0.013]	-0.116* [0.016]
δ_{KT}	0.002* [0.0005]	0.007* [0.003]	0.008* [0.0007]	0.005* [0.0004]	0.006* [0.0007]
β_L	0.211* [0.006]	0.389* [0.005]	0.348* [0.005]	0.347* [0.008]	0.529* [0.002]
β_{LL}	0.102* [0.021]	0.153* [0.023]	0.109* [0.014]	0.083* [0.015]	0.002 [0.004]
ρ_{LY}	-0.041* [0.014]	-0.0439 [0.026]	0.029 [0.028]	0.072* [0.021]	0.115* [0.016]
δ_{LT}	-0.003* [0.0008]	-0.006* [0.0003]	-0.007* [0.0004]	-0.008* [0.0003]	-0.006* [0.0007]
k	1	2	1	1	1

Diagnostic Tests

Share Equations	S_K	S_L	S_K	S_L	S_K	S_L	S_K	S_L	S_K	S_L
JB	0.90 (0.64)	0.31 (0.86)	1.94 (0.38)	1.67 (0.43)	0.95 (0.62)	0.43 (0.80)	1.04 (0.59)	0.91 (0.63)	0.36 (0.83)	1.66 (0.43)
RESET test	0.45 (0.73)	0.40 (0.79)	2.91 (0.08)	2.36 (0.11)	0.57 (0.81)	0.11 (0.99)	0.16 (0.99)	2.99 (0.08)	2.36 (0.11)	2.69 (0.09)

Notes: *Robust standard errors are derived using the Newey and West (1987) procedure. Figures in brackets denote standard errors. The optimal lag and lead structure (k) for DSUR was selected by minimizing the Schwarz criterion. (*) indicates statistical significance at the 5% significance level. Figures in parentheses represent asymptotic p-values associated with the tests. JB denotes the Jarque –Bera normality test of errors. The RESET tests the null hypothesis that there is no functional form misspecification.*

Table 4. Estimated long-run own and partial elasticities of substitution

Country	E_{KK}^{LR}	E_{LL}^{LR}	E_{MM}^{LR}	σ_{KL}^{LR}	σ_{KM}^{LR}	σ_{LM}^{LR}
Greece	-0.35* (0.05)	-0.32* (0.09)	-0.37* (0.08)	0.75* (0.04)	0.91* (0.05)	-0.59 (0.40)
France	-0.48* (0.006)	-0.22* (0.05)	-0.14 (0.09)	0.77* (0.03)	1.07* (0.07)	-0.83** (0.30)
Italy	-0.49* (0.009)	-0.34* (0.04)	-0.38** (0.04)	0.86* (0.03)	1.08* (0.03)	-0.91* (0.17)
Spain	-0.52* (0.007)	-0.40* (0.04)	-0.41** (0.07)	0.92* (0.02)	1.11* (0.06)	-0.32 (0.23)
UK	-3.81 (0.011)	-0.53* (0.007)	-5.56* (0.004)	1.01* (0.02)	1.88* (0.003)	1.01* (0.006)

***Notes**: Figures in parentheses are asymptotic standard errors. A (*) indicates statistical significance at the 5% significance level* .

Table 5 : Comparison of Allen Partial Elasticities with previous studies

	Italy		Spain		UK	
	(1)	(2)	(1)	(3)	(1)	(4)
σ_{KL}	0.78	0.83	0.90	1.13	1.34	1.04
σ_{KM}	1.02	0.30	0.86	0.29	1.32	0.92
σ_{LM}	-0.38	1.09	-0.18	0.94	1.01	1.02
Data	Time Series 1953-1977		Time Series 1971-1990		Time Series 1953-1984	

(1): This study. (2) Apostolakis [1981]. (3) Truett and Truett [2000] (4) Apostolakis [1988].

Table 6. Exogeneity Tests

Country	$\Delta\log(p_K / p_M)$	$\Delta\log(p_L / p_M)$	$\Delta\log(Y)$
Greece	0.05 [0.97]	3.29 [0.20]	0.71 [0.70]
France	0.71 [0.70]	6.96** [0.03]	0.90 [0.63]
Italy	2.13 [0.34]	1.70 [0.43]	0.08 [0.96]
Spain	7.69** [0.02]	0.27 [0.87]	0.44 [0.80]
UK	2.06 [0.35]	0.94 [0.62]	5.62* [0.06]

***Notes**: All tests are* χ^2 *statistics. Figures in brackets represent asymptotic p-values associated with the tests. ** and * rejection of the null hypothesis at the 5% level and 10% level of statistical significance respectively.*

Table 7. Tests of restrictions in the short-run

Restrictions	*Greece*	*France*	*Italy*	*Spain*	*UK*
$\alpha_K = \alpha_L$	1.49 [0.15]	0.30 [0.50]	1.25 [0.32]	0.09 [0.70]	1.12 [0.24]
$\varphi_K = \varphi_L$	1.52 [0.18]	0.008 [0.80]	0.54 [0.42]	0.07 [0.65]	0.09 [0.74]

Notes: *Figures in brackets represent p-values associated with* χ^2 *tests of the restrictions.*

Table 8. GMM estimates of error correction models

$$\Delta S_{Kt} = \alpha \Delta S_{K,t-1} + \sum_{i=0}^{q} \gamma_{KK,i} \Delta \log(p_K / p_M)_{t-i} + \sum_{i=0}^{q} \gamma_{KL,i} \Delta \log(p_L / p_M)_{t-i} + \sum_{i=0}^{q} \theta_{YK,i} \Delta \log Y_{t-i} + \phi EC_{K,t-1}$$

$$\Delta S_{Lt} = \alpha \Delta S_{L,t-1} + \sum_{i=0}^{q} \gamma_{KL,i} \Delta \log(p_K / p_M)_{t-i} + \sum_{i=0}^{q} \gamma_{LL,i} \Delta \log(p_L / p_M)_{t-i} + \sum_{i=0}^{q} \theta_{Y,iL} \Delta \log Y_{t-i} + \phi EC_{L,t-1}$$

Parameter	*Greece*	*France*	*Italy*	*Spain*	*UK*
α	-0.065 [0.07]	-0.42** [0.058]	0.028 [0.050]	0.09* [0.087]	0.005 [0.040]
γ_{KK0}	0.033** [0.011]	0.004** [0.002]	0.033** [0.009]	0.03** [0.007]	-0.677** [0.171]
γ_{KL0}	-0.036** [0.010]	0.007* [0.004]	-0.030** [0.001]	-0.017** [0.006]	0.051** [0.009]
θ_{YK0}	-0.039 [0.068]	-0.40** [0.05]	-0.270** [0.061]	-0.062* [0.045]	-0.113** [0.032]
ϕ	-0.283* [0.150]	-0.010** [0.005]	-0.045** [0.021]	-0.175** [0.064]	-0.191** [0.050]
γ_{LL0}	0.177** [0.019]	0.191** [0.018]	-0.108* [0.061]	-0.062 [0.076]	-0.053** [0.011]
θ_{YL0}	4.83×10^{-4} [0.050]	0.04** [0.009]	-0.042 [0.070]	-0.005 [0.065]	0.185** [0.029]
System R^2	0.88	0.97	0.93	0.90	0.99

Diagnostic Tests

Share Equations	S_K	S_L	S_K	S_L	S_K	S_L	S_K	S_L	S_K	S_L
JB	0.48 (0.78)	1.06 (0.56)	0.21 (0.86)	0.40 (0.76)	0.40 [(0.81)	0.47 (0.71)	1.42 (0.48)	1.95 (0.37)	1.20 (0.35)	0.90 (0.40)
RESET Test	0.78 (0.46)	1.03 (0.37)	1.02 (0.37)	1.82 (0.23)	0.77 (0.43)	0.24 (0.77)	0.76 (0.49)	1.72 (0.19)	3.02 (0.06)	0.62 (0.54)

Notes: *Figures in brackets are standard errors. ** and * indicate statistical significance at 5% and 10% significance level respectively. The optimal lag was selected by minimizing the Schwarz criterion. Figures in brackets denote standard errors. Figures in parentheses represent asymptotic p-values associated with the tests. JB denotes the Jarque –Bera normality test of errors. The RESET tests the null hypothesis that there is functional form misspecification.*

Table 9. Estimated Short-Run Own and Partial Elasticites of Substitution

Country	E_{KK}^{SR}	E_{LL}^{SR}	E_{MM}^{SR}	σ_{KL}^{SR}	σ_{KM}^{SR}	σ_{LM}^{SR}
Greece	-0.35* [0.01]	-0.28* [0.10]	-0.38* [0.16]	0.74* [0.07]	1.02* [0.13]	-0.81 [0.60]
France	-0.36* [0.01]	-0.23* [0.04]	-0.27 [0.20]	0.68* [0.04]	0.74* [0.09]	-0.11 [0.25]
Italy	-0.46* [0.02]	-0.35* [0.01]	-0.36* [0.10]	0.76* [0.03]	0.99* [0.15]	-0.37* [0.17]
Spain	-0.47* [0.02]	-0.40* [0.05]	-0.35* [0.12]	0.90* [0.02]	0.86* [0.15]	-0.18 [0.38]
UK	-2.79* [0.58]	-0.52* [0.07]	-3.77* [0.89]	1.34* [0.05]	1.32* [2.82]	1.01* [0.06]

***Notes**: Figures in brackets are asymptotic standard errors. A (*) indicates significance at the 5% level of statistical significance.*

Appendix

The cost function approach developed in section 2 requires data on the prices and quantities of each input as well as the level of output to final demand. All data required for this study were extracted from the European Union's AMECO database (Annual Macro Economic Data Base DG2) from 1960 through 1998.

Following Aw and Roberts [1985] output (Y) is measured as gross domestic product (GDP) plus the value imports and minus the non –factor costs:

$$Y = GDP + P_M M - T_{ind} + S$$

where T_{ind}, $P_M M$ and S are indirect taxes, value of imports and subsidies respectively.

The quantity of output is constructed as current value output divided by the output deflator.

The implicit price index of capital input is constructed by dividing capital expenditures by net capital stock. Capital expenditures is measured as GDP minus the compensation of labor.

The price series of labor was calculated as the ratio of total compensation of employees to total employment.

The price of import was constructed by dividing the current value of imports by the value of imports at constant prices.

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Endnotes

[1] In a previous version of this paper we have reported 3SLS results. However, additional LM tests for autocorrelation rejected the null in most cases so we report hetorescedasticity and autocorrelated consistent (HAC) standard errors derived from GMM. 3SLS cannot allow that. Based on 3SLS standard errors the error correction term would have been statistical insignificant (see ϕ coefficients in Table 8). Based on GMM HAC standard errors this is, however, no longer the case.