On the Determinants and Effect of Employment: An Empirical Assessment of Tourism and Innovation

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Submitted to the
Institute of Graduate Studies and Research
in partial fulfillment of the requirements for the degree of

Doctor of Philosophy
in
Economics

Eastern Mediterranean University
September 2020
Gazimağusa, North Cyprus

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ABSTRACT

The primary objective of this theses is to determine the determinants and effect of employment with particular regards to innovation and tourism. To this end, the thesis is divided into two self-contained sections.

In the first section, the innovation employment nexus is analysed in a panel of 8 Asian economies within the framework of a panel cointegration methodology. Pooled mean group, mean group and the dynamic fixed effects estimators were employed to obtain the short-run values and the long run equilibrium values within a linear and non-linear specification. While the linear specification produced mixed results, the non-linear specification indicated a U-shaped non-linear relationship between r&d and employment with a local minimum at about the 75th percentile of the r&d data range.

In the second section panel cointegration methodologies were employed to ascertain the impact of innovation on sectoral employment. FMOLS estimation results show that while r&d is employment creating in the services and high-tech manufacturing sectors, it is however employment constraining in the low-tech manufacturing sector.

The third section examines how employment affects demand for tourism in the short and long run, controlling for the effects of income and relative prices within a panel of 32 Organisation for Economic Co-operation and Development (OECD) countries throughout the 1995–2016 period. Because of this, second-generation panel unit root tests, panel cointegration tests and panel data estimation techniques are employed. Results indicate that while employment has a positive association with outbound tourism in the short-run, its positive effect on outbound tourism in the long-run is however insignificant. Causality results uncover causality flowing from income to employment with a feedback, uni-directional causality flowing from income to outbound tourism and causality flowing from relative prices to outbound tourism with a feedback.

Keywords: Innovation, Employment, Panel Cointegration, Cross-sectional Dependence, Causality.

ÖZ

Bu tezin temel amacı, özellikle inovasyon ve turizm açısından istihdamın belirleyicilerini ve etkisini belirlemektir. Bu amaçla, tez iki bağımsız bölüme ayrılmıştır.

İlk bölümde, inovasyon istihdamı bağlantısı, bir eşbütünleşme yöntemi çerçevesinde 8 Asya ekonomisinden oluşan bir panelde analiz edilmektedir. Havuzlanmış ortalama grup, ortalama grup ve dinamik sabit etkiler tahmin ediciler, doğrusal ve doğrusal olmayan bir spesifikasyon dahilinde kısa dönem değerlerini ve uzun dönem denge değerlerini elde etmek için kullanılmıştır. Doğrusal belirtim karışık sonuçlar üretirken, doğrusal olmayan belirtim, ar-ge ve istihdam arasında, ar-ge veri aralığının yaklaşık 75. yüzdelik diliminde yerel bir minimum ile U şeklinde doğrusal olmayan bir ilişki gösterdi.

İkinci bölümde inovasyonun sektörel istihdam üzerindeki etkisini belirlemek için panel eşbütünleşme metodolojileri kullanıldı. FMOLS tahmin sonuçları, Ar-Ge, hizmetler ve yüksek teknolojili imalat sektörlerinde istihdam yaratırken, düşük teknolojili imalat sektöründe istihdamın kısıtlandığını göstermektedir.

Üçüncü bölüm, 1995-2016 dönemi boyunca 32 Ekonomik İşbirliği ve Kalkınma Örgütü (OECD) ülkesinden oluşan bir panelde gelirin ve nispi fiyatların etkilerini kontrol ederek, istihdamın kısa ve uzun vadede turizm talebini nasıl etkilediğini incelemektedir. Bu nedenle ikinci nesil panel birim kök testleri, panel eşbütünleşme testleri ve panel veri tahmin teknikleri kullanılmaktadır. Sonuçlar, istihdamın kısa vadede giden turizm ile olumlu bir ilişkisi olsa da, uzun vadede giden turizm üzerindeki olumlu etkisinin önemsiz olduğunu göstermektedir. Nedensellik sonuçları, geri bildirim ile gelirden istihdama akan nedenselliği, gelirden giden turizme akan tek yönlü nedenselliği ve bir geri bildirimle göreli fiyatlardan giden turizme akan nedenselliği ortaya çıkarır.

Anahtar Kelimeler: Yenilik, Istihdam, Panel Eşbütünleşmesi, Kesitsel Bağımlılık, Nedensellik.

DEDICATION

To My Family

ACKNOWLEDGEMENT

I want to give a special note of thanks to my supervisor Prof. Dr. Mehmet Balcılar who spared me a lot of his valuable time throughout the duration of this research. His immense understanding, patience and expertise in the field were priceless attributes which I am exceedingly grateful for. I want to thank all members of staff of the department of economics in particular and the faculty of business and economics in general who in no small way contributed to the fulfilment of this program. To all my friends, colleagues and course mates who are too numerous to mention, I thank you all for the numerous ways in which you all have contributed to the success of this program and the fond memories we share. I want to also thank my parents whose guidance and love helped me throughout this journey, my love for them is immeasurable. I also want to thank my brother and sisters who were also there for me every step of the way. I could never thank them enough for their love, patience and understanding.

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Chapter 1

INTRODUCTION

Employment is a lagging macroeconomic indicator and as such tends to improve or deteriorate due to economic expansions and contractions respectively. Thus, it is expected that the state of the economy would reflect on the level of employment in any economy because labour is an essential requirement of production. Technical progress which is embodied in novel production equipment or disembodied from productivity gains from new products which are unrelated to productivity can significantly alter the production process as well as the skill requirement and intensity of labour. Changes in employment level would thus have further implication on the consumption patterns of individuals. The first primary objective of the present thesis is to analyse empirically the extent to which innovation affects employment in a panel of eight Asian countries. The thesis would analyse the non-linear employment effect of aggregate r&d expenditures in eight Asian countries. Moving further to the second main objective, the thesis would analyse the job creating impact of aggregate r&d expenditures on the sectoral employment levels of 12 European countries. The third main objective of this thesis is to ascertain how employment affects the consumption of tourism in a panel of 32 OECD countries. This would give a holistic perspective of how to initiate demand and supply side policies related to employment taking tourism demand and innovation into perspective.

The welfare consequences of innovation as regards to job displacement vis-a-vis creation has been a major concern to economists since the advent of the industrial revolution. While Ricardo (1951) emphasized the fear of the working class as to the potential of their services being replaced by machines, a viewpoint which was previously physically validated by the early 19th century Luddite riots, Marx (1961) on the other hand postulated the “compensation theory” which addressed the counterbalancing effect of excess capital made available from the marginal productivity gains of labour saving process innovation on initial job losses. Both Marx and Ricardo held the same view points as regards to the initial negative welfare effects of process innovation on the working class. Mokyr et al (2015) were also of the view that these concerns were much more emphasized at significant business cycle episodes that reflected technological change and/or economic recessions. Innovation greatly stimulates economic growth and development but the net effect on job creation remains largely unclear. Two sections of this thesis intend to isolate at the macroeconomic level—how process innovation as well as product innovation induces job creation in eight Asian economies.

Following the employment effects of innovation, the thesis further attempts to isolate the outbound tourism demand effects of employment in 32 OECD countries while controlling for the individual effects of price and income. The employment effect of outbound travel was empirically investigated by Chi (2016) for US tourists destined for Hawaii. This thesis intends to broaden the demographic scope of Chi (2016) by empirically determining how aggregate outbound tourism expenditures in OECD countries responds to the movement of employment rates. This section of the thesis has important policy implications for both potential destination countries and departure countries accordingly. For potential destination countries, policy frameworks can be synchronized with employment expectations in OECD countries in order to formulate better marketing strategies which takes into account the prices, income and employment of OECD countries. The importance of employing OECD data lies in the fact that these countries represent the most advanced economies in the world and have a large share of the worldwide tourism market as regards to expenditures. Research implications for OECD countries naturally follows a better ability to predict potential tourism consumption based on employment, relative prices and the state of the economy. This would enable OECD countries to better tax these consumption activities.

Subsequent sections of this theses include a general literature review in chapter 2, in chapter 3 the presents the study on innovation and employment in Asian economies, chapter 4 presents the study on the inter-sectoral employment effects of aggregate r&d expenditures, chapter 5 presents the study on employment and outbound tourism demand in OECD countries while chapter 6 concludes.

Chapter 2

LITERATURE REVIEW

2.1 Innovation and Employment

The literature on R&D and employment stretches all the way from the 90's, there is a large literature on the analysis of the nexus between changes in employment due to innovation. Initial research, due to data availability were mostly cross-sectional in nature. Zimmerman (1991), employing data from 16 industries in Germany, came to the conclusion that innovation shocks was one of the critical factors enabling the German employment decrease in the 80s. His research for the most part uncovers a negative innovation impact. However, his definition of innovation refers to a question explicitly inquiring whether or not labour-saving technological progress was implemented. Brouwer, Kleinke and Reijnen (1993), employing data from 859 Dutch manufacturing firms, uncovered a total inverse connection between aggregate R&D expenditures and employment while the reverse scenario emerges when only the case of product innovation is considered. Blanchflower and Burgess (1998) uncovers a direct relationship amongst process innovation (measured with a dummy) and the growth in employment employing innovation survey from the UK in 1990 and Australia in 1998/1990. More contemporary research has advantage of new available datasets with panel structures and have employed panel data econometric methodologies that mutually consider time and cross-sectional inconsistency. Piva and Vivarelli (2005), applied the system GMM method to explore whether the positive effect technological change has on job creation is still valid in conditions where intermediate technologies are implemented mainly through non-R&D expenditures in Italian manufacturing firms based on different survey by MCC (Mediocredito Centrale Investment Bank) from 1992-1997. They find a significant – although slight positive relationship amongst gross innovation investment and employment at firm level which was robust to time, industry, size of firm and geographical fixed effects.

Evangelista and Savona (2002), use an innovation survey undertaken in Italy by the National Statistical Office (ISTAT) for the Italian service industries in (1993-1995), to determine the employment effects of innovation and uncover a direct positive impact in the most innovative and knowledge intensive service sectors in the case of financial-related sectors. Lachenmaier and Rottmann (2011), use panel dataset from IFO innovation survey for manufacturing in Germany in (1982-2002) to ascertain the role of product and process innovation, employing a generalized method of moments (GMM) framework. They find a positive relationship between persistent process innovation activities and employment of current and lagged product and process innovation, with the size of the impact rising with the time lag, while, a smaller effect is found for persistent product innovation. Antonucci and Pianta (2002), use European Innovation Survey (CIS II) for 8 European countries to estimate effect of innovation on employment for 1994-1999 depending on innovative and other control variables from 1994-1996. They find a negative impact of innovation on European manufacturing employment. Evangelista and Savona (2003), use CISII for the period 1993-1995, for the Italian firm to estimate direct effect of innovation strategy perused by firms, across industries and firm's skill intensity. They find a negative overall effect of innovation on employment. In Italian services sectors heavy job losses were found in the largest firm, among low skill workers, in sectors heavy users of ICTs, in capital intensive and financial-related one, net job creation emerged in smaller firms and in technology-oriented activities. More recent study further explored the displacement or compensation mechanisms due to different types of innovation.

Bogliacino and Vivarelli (2012), running GMM-SYS and a dummy variable corrected least square analysis on manufacturing and services sectors for 15 European countries within the time span 1996-2005, found that R&D expenditure shows a positive employment inducing effect in the manufacturing (high-Tech) and services sector, while not relevant in the (Low-tech) manufacturing sectors. They present evidence for a positive employment effect of R&D expenditures, particular concentrated in high-Tech segment. Van Roy et.al (2015) model a labour demand equation augmented with (lagged) innovation by GMM method by using patent data in a sample of European patents, they find positive impact of patent for firms in high-Tech manufacturing sector but not for firms in low-Tech manufacturing sectors and service sector. Mastrostefano and Pianta (2009) empirically isolate the employment inducing effect of innovation at the industry level with data for 10 European countries, by employing dual EU innovation surveys – CIS2 (Community Innovation Survey) (1994–1996) and CIS3 (1998–2000), running the OLS and GLS model found a positive employment impact of innovation which is obvious in the high-tech manufacturing sectors.

Overall, current studies, specifically those based on consistent panel data analyses, offer a comprehensive account of the job creation impact of innovation as they overcome their possible job displacement effects. The most recent panel investigation tends to support that in the sectoral dimension, labour-friendly impact is generally limited to high-tech sectors and services sectors however does not hold in the low-tech manufacturing sectors, particularly when R&D or product innovation or their combination are utilized to proxy technical change. It is important to note that most of the studies used GMM-SYS methodology but our model will utilize the FMOLS methodology in order to estimate the long-run coefficients using the cointegrating relationship.

Earlier empirical studies on the employment innovation nexus began with the study by Entorf and Pohlmeier (1990), they uncover a direct relationship between product innovation and employment employing data for over 2,000 firms of West German origin, in the single time period of 1984. This relationship was further verified by Smolny (1998) also employing data for over 2,000 Manufacturing firms based in the west of Germany during the period 1980-1982. Later studies like Harrison et al employing manufacturing and services data, validate the existence of a compensating mechanism inherent in the innovation employment nexus over a two-year period across 20,000 companies within France, Germany, Spain and the UK, their analysis supports the job displacement effect of process innovation whilst confirming the counterbalancing effect of increased demand for older products.

Within Asia, Mehta (2016) analyze the impact of product innovation on four different industries of the Indian manufacturing sectors ranging from the low to the high technology industries within the periods 2000 to 2014 and validated a positive relationship between product innovation and employment growth. Kwon et al (2015) verify the effects of process and product innovation on 532 manufacturing firms in South Korea and uncover a positive connection between product innovation and the ability of the firms to create new employment, however the reverse effect was obtained for process innovation.

Ciriaci et al. (2016) employed a semi-parametric quantile regression approach for a panel of 3304 Spanish firms within the periods 2002- 2009. Their results indicated that episodic high employment growth rates were better sustained by younger, smaller and more innovative firms than non-innovative firms.

A number of similar studies (Dachs et al. 2017; Yu and Yu 2017; Evangelista and Savona 2002; Lanchenmaer and Rottmann, 2007; 2011 amongst others) also uncover evidence for the job creating effect of innovation although in varying degrees across different economies, however most of these studies employ micro-economic data in their estimations and thus their results may not be an adequate representation of their respective macro-economies.

2.2 Determinants of Outbound Tourism Demand

One of the most significant incentives to undertake a holiday trip in the outbound tourism demand literature is personal income which has been measured by GDP (Halcioglou, 2010), GDP per capita (Balli et al., 2019) or average wages (Chi, 2016). By employing a general to specific methodology within the framework of an error correction model, Song et al. (2000) estimates income and price elasticities for 11 travel destinations and discovers that while overseas travel from the United Kingdom is highly income elastic, the own-price elasticity of demand for UK outbound travels is less than unity. This suggests that while outbound tourism is a luxury in the United Kingdom, price increments from destination countries would generate net revenue effects for these countries. The Chinese demand for tourism in Thailand is also investigated in a study by Untong et al (2015). It is discovered through their empirical analysis that tourism to Thailand is perceived as a luxury by the Chinese owing to the greater than unity coefficient obtained for Chinese real GDP. Also, obtained price elasticities show that China is quite sensitive to destination price appreciation due to a greater than unity coefficient on the real exchange rate adjusted relative prices of Thailand. In the same vein, Kim et al. (2012) and Fereidouni et al. (2017) have shown that earnings from capital are seen to significantly affect the decision to take a holiday trip to other parts of the world. They both incorporate housing price indices in their demand model to analyse if wealth effects from real estate have any significant impact on outbound tourism demand in South Korea and Malaysia, respectively. They discover that wealth effects from real estate significantly increase outbound travel demand from the two countries. Seetaram (2012) estimates Australia's outbound travel demand elasticities for 47 travel destinations and uncovers a significant, above unity income effect and a positive net migration effect. The model shows that other factors besides relative prices and income such as prior information about destination countries from migrants and visits to their countries of origin may also be major incentives to undertake holiday trips.

The macroeconomic conditions of destination countries relative to departure countries can also contribute to the tendency to travel or the propensity to spend at destinations. One variable that is widely used to capture this effect is the exchange rate adjusted relative prices. Zhang et al. (2012) employ descriptive and factor analyses to identify the preferences and factors influencing the outbound travel demand residents in Shanghai during the global financial crisis (GFC). They discover that the GFC-induced macroeconomic effects such as devaluation of destination currency and the attendant relative price drops significantly affect Shanghai residents' outbound travel decisions in the positive direction. Cort´es-Jim´enez et al. (2009) employing the error correction based linear almost ideal demand system uncover different income and price elasticities of Italian demand for four different tourist destinations. The short-run elasticities from their empirical investigation uncovered more significant coefficient estimates than the long-run estimates. Going backwards, Crompton (1979) develops the 'push and pull' factors model that influences the decision-making process of tourists as regards to socio-psychological motivations and travel destination attractions. He attributes the push factors to innate factors within and around the individual which instigate a desire to travel, while the pull factors are qualities possessed by the travel destination which attract the individual. The model remains one of the major influences in the construction of inbound and outbound tourism demand models. Just like some of the above-mentioned studies (Fereidouni et al., 2017; Kim et al., 2012; Seetaram, 2012), our study abstracts away from pull factors to cover a broader scope within the framework of outbound travel demand. A few other studies using panel data estimation techniques employ different specifications of the gravity model to capture the long-run equilibrium relationship between tourism and its long-run demand determinants (Brun et al., 2005; Eryigˇit et al., 2010; Khadaroo and Seetanah, 2008; Yazdi and Khanalizadeh, 2017). This allows for the incorporation of both push and pull factors in the model by analysing tourism inflows from multiple countries to a single country or outflows from a single country to multiple destinations. Within the context of the present study, incorporating push and pull factors entails controlling for the demand elasticities of both destination and departure countries.

This study, however, follows the approach of Halicioglu (2010), which employs the aggregate outbound tourist flows, relative price effects and the GDP data of Turkey to estimate its outbound tourism demand elasticities. Halicioglu (2010) uncovered a significantly elastic and above unity income effect and a negative relative price effect, which implies that tourism is a luxury in Turkey. The study indicates that causality flows only from income to outbound tourism demand in both the long-run and the short-run indicating that income is an important variable in the prediction of outbound tourist flows in Turkey. Halicioglu (2010) differs from the present study in the sense that his study employs time-series cointegration methodologies to estimate tourism demand elasticities for a single country. The present study, however, employs heterogeneous panel cointegration and estimation techniques to estimate the long-run tourism demand elasticities for a panel of 32 OECD departure countries while also controlling for employment effects. This has the advantage of controlling for the outbound tourism demand determinants of multiple departure countries, which broadens the geographical scope of the present research. In the same vein, Dogru and Sirakaya-Turk (2018) employ panel cointegration techniques to test the outbound tourism demand elasticities for major Turkish tourist destinations and uncover a significantly positive income effect which is less than unity. The result implies that outbound tourism to these destinations is a nonluxury item for Turkish tourists. This might be because unlike the Halicioglu (2010) study which employs aggregate data, their data are limited to 12 Turkish tourist destinations. This can suggest that different data aggregation and/or measurement approaches can produce diverse results even for the same study location. It can also be as a result of business cycle volatility in the departure (destination) country, which results in changing demand elasticities across different business cycle phases as broadly elucidated by Smeral (2012), Smeral and Song (2015) and Gunter and Smeral (2016, 2017). Smeral (2012) exploits the advantages of a time-varying parameter model to isolate the asymmetric income and price effects inherent in outbound tourism demand for various source markets. The study uncovers that tourism import elasticities are heterogeneous across different business cycle phases. Other studies (Agnew and Palutikof, 2006; Mckercher and Hui, 2004; Moore, 2010; Yang and Wu, 2014) have employed various methodologies to estimate outbound tourism demand elasticities while controlling for variables of study specific interest. However, only a few of these studies (for instance, Chi, 2016) control for employment effects. In the study by Chi (2016), employment and wage effects are incorporated in a model to explain tourism demand to Hawaii from the US mainland. He discovers that tourism demand from the United States responds positively to increased employment and wage levels. Our study goes a step further to see if this relationship holds within a broader set of geographical locations by controlling for outbound tourism demand effects of employment in a panel of 32 OECD countries. Our choice of OECD countries is based on the earlier observation that expenditures from the OECD tourism market makes up about half of the global tourism market. Also, most of the studies reviewed are limited in their scope of departure countries which capture a very limited fraction of the international tourism market and thus portends very limited implications for potential inbound tourism demand to other destination countries.

The fact that employment generates disposable income would make the level of employment in any country a valid determinant of consumption. However, increasing levels of employment may not translate to increased consumption patterns especially at per capita levels due to dissimilar working conditions and employment remuneration across different job endeavours. A scenario may arise wherein the level of employment is a determinant of outbound tourism demand in the long-run (Chi, 2016) with the implication of improved working conditions in the long run. This would imply that jobs with wages sufficient enough to offset the cost of travel trips are being created more than jobs with less than sufficient wages. Employment might have further impact on outbound tourism demand beyond its income-generating property because it builds confidence in consumers. We also control for tourism prices by including the effective exchange rates adjusted relative prices as proxy. Furthermore, our study takes into consideration the time-series properties of the variables employed in the demand equations as neglecting this very important factor can lead to models with spurious long-run coefficient estimates. Cross-country heterogeneity and dynamics have also been neglected in quite a number of studies that employ panel data. This also has the undesired effect of drawing inferences that are potentially misleading (Haque et al., 1999; Pesaran and Smith, 1995). Additionally, there are instances wherein countries across the panels are all affected in varying degrees by unobserved common shocks which can ultimately lead to cross-sectional dependence across countries. If the effects of common factors are not partialled out, the efficiency gains of panel data techniques over single equation estimation techniques can be significantly reduced if not entirely lost (Pesaran, 2004). It is to this end that we employ dynamic heterogeneous panel cointegration techniques to ascertain the long-run equilibrium and causal relationship between tourism demand and employment using a dataset of selected OECD economies. To the best of the author's knowledge, this will be the first of such studies and as such is intended to fill the aforementioned gaps in the literature.

Chapter 3

INNOVATION AND EMPLOYMENT IN ASIAN COUNTRIES

3.1 Introduction

The literature on Innovation and employment is riddled with diverse opinions emanating from different viewpoints. Most researchers emphasize that the labour saving impact of R&D investments are usually counter balanced by job creating technological change (Vivarelli,2012). However, the lag period in which this adjustment occurs may be characterized by a painful short-run adaptation process (Oberdabernig,2016). Schumpeter (1942) Terms this process "creative destruction" whereby older methods of production are rendered obsolete via the introduction of newer more refined methods. This may have the net effect of replacing unskilled labour with skilled employment (Damijan et al, 2014; Ugur et al,2016). Calvino and Virgillito (2016) emphasize two theoretical perspectives on the innovation-employment nexus. The first equilibrium perspective assumes an adjustment process whereby increased innovation instigates higher employment via increased output and reduced wages, however the second disequilibrium perspective assumes inherent complexities in technical progress and as such, renders the employment-innovation nexus ambiguous and intractable.

Studies analyzing product (process) innovation-employment nexus were mostly undertaken with firm-level data and were usually focused on Latin American countries. These studies mostly reinforce the employment creating effect of product innovation (Castillo et al 2014; Crespi and Tacsir,2013; Álvarez et al 2011; Aboal et al,2015 amongst others) Most of these studies utilize different proxies for process and product innovation, while the majority of them uncover a positive relationship between product innovation and employment, the relationship between process innovation and unemployment was quite disparate across studies (Calvino and Virgillito, 2016).

Most studies of an empirical nature underline the need to differentiate product innovation from process innovation. Product innovation generates more employment via the introduction of new products which would increase demand resulting in the need for more workers in the innovating firm(s) in order to maintain equilibrium with the demand and supply side. Another scenario may arise wherein the firm having gained monopoly power and in a bid to maximize its profit margin may not absorb more workers or may not need to if the creation of its new products necessitates the utilization of less labour, especially if this innovation was a replacement of older preexisting products thus resulting in lesser output and lesser employment. Process innovation on the other hand may lead to the loss of jobs because the increased marginal productivity of labour that is a direct consequence of this innovation may lead to the production of the same level of output with less labour. This may however go both ways as the marginal productivity gains may be reflected in the price of the products and thus lesser product prices would lead to higher product demand which would invariably translate to higher labour demand. (see Vivarelli,2014; Lachenmaier and Rottman, 2007).

Dosi (1984) points out that the terms product innovation and process innovation are interchangeable across different sectors even for an individual product. So, in essence, the concept of process innovation being employment destroying and product innovation being employment inducing may be all too simplistic (Pianta,2005). The net effect of process and product innovation remains ambiguous at best (Vivarelli,2012). The r&d variable employed in this study incorporates both public and private r&d expenditure as such would give a more holistic account of its possible job creating effects. The firm level r&d expenditure employed in previous studies however do not also include public r&d expenditure. In disentangling the latent effect of product and process innovation on unemployment there arises a need to isolate to a greater detail the long-run dynamic relationship between these variables at a macroeconomic level. This study attempts to fill this gap. Also, attention has not really been paid to Asian economies in the literature, as a result this study's contribution to the literature comes in three folds. Firstly, by employing panel cointegration and estimation approaches the long run and short-run equilibrium parameters of the r&d expenditures/ employment relationship may be ascertained. By utilizing the pooled mean group (PMG) estimation of Pesaran et al (1997) an intermediary of the mean group estimator (MG) of Pesaran and Smith (1995) and the dynamic fixed effects (DFE), heterogeneous short-run dynamics and long-run slope homogeneity can both simultaneously be controlled for. The MG estimates are consistent under both long run slope homogeneity and heterogeneity while the PMG estimates are efficient under long run homogeneity and inconsistent under long run heterogeneity. Long run slope homogeneity is validated empirically through the Hausman test. Secondly the macro-economic dataset employed in the estimation would give a more holistic account of the r&d job creating effect in the panel. Lastly by incorporating non-linearities in the long run cointegrating space, potential non-linear threshold effects in the R&D-employment relationship may properly be accounted for.

3.2 Data and Methodology

Data for yearly, employment, Research and development expenditure (% of GDP), gross fixed capital formation (GFCF) and GDP which was used as a proxy for production was gathered from the World Bank's World Development Indicators. Data for monthly, minimum wage was gathered from the International Labour Organization. They were selected based on data availability for the 8 Asian countries under study, within the period 1997- 2014. The GDP and GFCF variables were measured at constant 2010 USD prices, the consumer price index was employed to deflate the monthly nominal minimum wage to real values, after which the real minimum wage was converted to dollar amounts by dividing with the relevant country currency/US dollar exchange rate. Before proceeding to the empirical analysis, all independent variables were transformed to natural logarithms except for r&d and employment that were measured in percentages of GDP and normalized on country population respectively. The selected 8 Asian countries are Armenia, China, Israel, Kazakhstan, Korea, Russia, Japan and Turkey.

3.2.1 The Employment Demand Equation

The equation for employment demand for the 8 Asian countries can be specified implicitly as;

$l=f(\mathrm{l}gi,\mathrm{l}w,r\&d,\mathrm{l}gdp)$

Where l stands for employment to population ratio, l*gi* stands for fixed capital formation (gross), l*w* stands for minimum wage, *r&d* stand for research and development expenditure and l*gdp* stands for production. The stochastic version of the model (see for similar approach, Lanchenmaer and Rottmann, 2011, Bogliacino and Vivarelli, 2012) for countries (i), over time (t) is specified as:

$$l_{it} = \beta_0 + \beta_1\, l\, w_{it} + \beta_2 r\&d_{it} + \beta_4 lgi_{it} \quad \beta_5 lgdp_{it} + v_{i,t} \tag{1}$$

$$l_{it} = \beta_0 + \beta_1\, l\, w_{it} + \beta_2 r\&d_{it} + \beta_3 r\&d_{it}^2 + \beta_4 lgi_{it} \quad \beta_5 lgdp_{it} + v_{i,t} \tag{2}$$

where v is the stochastic error term.

In model (1) the r&d employment relationship is assumed to be linear. However, in model (2) non-linearity is introduced by specifying the r&d variable in the quadratic form.

3.2.2 Panel unit Root Test

The mean reverting property of each variable mentioned in (1) was verified through unit root test: The Fisher-ADF tests of Maddala and Wu (1999) and Choi (2001), LLC test of Levin et al (2002) and IPS test of IM et al (2003) were all employed for the tests. These unit root analyses indicate the null hypothesis to be the presence of a unit root against the alternative of mean reversion. Two modes are employed for the unit root tests in levels and first differences. The most prevalent equation for the panel unit root test (Levin et al, 2002; Dicky and Fuller, 1979) is explicitly modeled as:

$$\Delta y_{it} = \alpha_i + \beta_i y_{it-1} + \sum_{j=1}^{pi} p_i \Delta y_{it-j} + e_{it} \tag{3}$$

Where p has to be selected in such a way that the residuals would be uncorrelated over time, Δy_{it} indicates y_{it} at first difference for country i, at time period $t = 1, \ldots, T$. Because this method (LLC) assumes homogeneity within the panel, β_i is thus homogeneous for all countries. The null hypothesis is that H_0: $\beta_i = 0$ against the alternative H_1: $\beta_i > 0$ for some i, which entails stationarity.

3.2.3 Panel Cointegration Test

A major advantage of panel data cointegration tests is that they have more efficiency gains than univariate approaches. The panel cointegration test of Pedroni (1999) was employed to ascertain the possibility of a long-run equilibrium relationship between employment and the independent variables. Recently, cointegration tests and panel model estimation procedures have elicited a lot of attention from researchers (see Larsson et al. 2001; Kao 1999; Harris and Sollis 2003). Dealing with the problem of non-homogeneous short-run dynamics and cross-country heterogeneity remains one of the most salient issues in those papers, for heterogeneity based issues appear to be predominant in analyses involving longitudinal data. The cointegration test of Pedroni (1999) was employed to diminish the heterogeneity problem.

Additionally, when processes with autoregressive (AR) errors are being modelled, parametric tests like the ADF-type tests are preferred because of its greater power as the regression model precisely identifies the AR terms, thus when the latent data generating process statistics are not known, it becomes more advantageous to use various testing procedures. The present study will employ both the between and within group cointegration testing procedures.

3.2.4 Econometric Estimation Method

In estimating equations (1) and (2) panel data methods which allow for both time-series and cross-section variation in all variables would be employed. Theoretically, cross-sectional estimation methods do not really identify any type of long-run connection between variables in the model because the time-series components of the data are not incorporated, which neglects potential efficiency gains. Estimating the model with panel data procedures based on; for instance pooled estimators is preferable, which could be applied to equation (1) and (2), however the aforementioned technique does not accommodate possibly heterogeneous dynamic adjustment around the long-run equilibrium relationship (Pesaran and Smith, 1995; Pesaran et al. 1999). In modelling the short-run dynamics the use of a more modified modelling approach is required. It is assumed that equation 1 and 2 denotes the long-run equilibrium relationship, but that the endogenous variable may diverge from its equilibrium path in the short-run due to the persistence of employment shocks overtime. By employing the mean group estimator (MG) (Pesaran and Smith,1995), the pooled mean group estimator (PMG) (Pesaran et al., 1999) and the dynamic fixed effects model (DFE), the parameters of equation 1 and 2 can be estimated. The three aforementioned estimation techniques can all be denoted by Eq (4). The DFE model happens to be the most restrictive of the trio by imposing homogeneity on the short-run and long-run parameters across countries, while allowing for intercept heterogeneity. Wrongly imposing homogeneity as aptly demonstrated by Pesaran and Smith (1995) even in large samples can bring about the inconsistency of pooled dynamic heterogeneous model estimates. The MG estimator lying at the other end of the spectrum as the least restrictive imposes no homogeneity and is obtained by averaging the individual country coefficients across the panel.

In between the MG and the DFE estimators is the PMG estimator which has homogeneity imposed on its long-run coefficients; $\theta_i = \theta$ for all i. However, the short-run parameters including the adjustment speed as well as error variances are allowed to differ across countries. The PMG model is appropriate regardless of the exogenous variables being I(1) or I(0). With this procedure the inconsistency problem is circumvented whilst the efficiency gains from pooled estimation are retained. The imposition of common long run coefficients is testable using a Hausman test and is conceivable given that the countries may face common technological and market conditions. Also plausible is the variability in short-run cross-country response, since the short run reactions depend on adjustment costs, financial considerations, and expectations. Following Pesaran et al. (1999), the panel analysis will be based on the unrestricted error correction ARDL (p, q) representation:

$$\Delta y_{it} = \Phi_i y_{i,t-1} + \beta_i^{'} x_{i,t-1} + \sum_{j=1}^{p-1} \lambda_{ij} \Delta y_{i,t-j} + \sum_{j=0}^{q-1} \gamma_{ij}^{'} \Delta x_{i,t-j} + \mu_i + u_{it} \quad (4)$$

In the equation above, y_{it} denotes the scalar endogenous variable, x_{it} is indicative of a $k\times 1$ vector of regressors for group *i*, the fixed effects is denoted by μ_i, Φ_i indicates a the lagged dependent variable scalar coefficient. The $k\times 1$ vector of parameters on explanatory variables is denoted by $\beta_i^{'}$, λ_{ij} denotes coefficients on lagged first-differenced terms of the dependent variable which are scalar by construction, and γ_{ij} are $k\times 1$ coefficient vectors on first-difference of explanatory variables and their lagged values. It's assumed that the disturbances u_{it} in the ARDL model are independently distributed across *i* and t, with constant means and variances $\delta_i^2 > 0$. It's Furthermore assumed that $\Phi_i < 0$ for all *i*, which validates the incidence of a long-run association between y_{it} and x_{it} defined by

$$y_{it} = \Theta_i^{'} x_{it} + \eta_{it} \quad i = 1, 2, \ldots, N; t = 1,2,\ldots,T \quad (5)$$

Where $\Theta_i^{'} = -{\beta_i^{'}}/{\Phi_i}$ is the $k\times 1$ vector of long-run coefficients, and η_{it} are stationary and mean reverting (including fixed effects). Equation (5) can be rewritten as:

$$\Delta y_{it} = \Phi_i \eta_{i,t-1} + \sum_{j=1}^{p-1} \lambda_{ij} \Delta y_{i,t-j} + \sum_{j=0}^{q-1} \gamma_{ij}^{'} \Delta x_{i.t-j} + \mu_i + u_{it} \quad (6)$$

In (6) $\eta_{i.t-1}$ is the error correction term while Φ_i is the error correction coefficient measuring the adjustment speed towards the long-run equilibrium.

The pooled maximum likelihood estimation is utilized to calibrate the group-specific short-run coefficients and the common long-run coefficients. These estimators are denoted by:

$$, \widehat{\Phi}_{PMG} = \frac{\sum_{i=1}^{N} \tilde{\Phi}_i}{N}, \widehat{\Phi}_{PMG} = \frac{\sum_{i=1}^{N} \tilde{\beta}_i}{N}, \tilde{\lambda}_{jPMG} = \frac{\sum_{i=1}^{N} \tilde{\lambda}_{ij}}{N}, \quad j = 1,\ldots,p\text{-}1$$

$$\hat{\delta}_{jPMG} = \frac{\sum_{i=1}^{N} \tilde{\delta}_{ij}}{N}, j=0, \ldots\ldots, q\text{-}1, \hat{\theta}_{PMG} = \tilde{\theta} \tag{7}$$

The MG estimator however (Pesaran and Smith 1995) controls for parameter heterogeneity and models the parameters (short-run and long-run) as:

$$, \widehat{\Phi}_{MG} = \frac{\sum_{i=1}^{N} \tilde{\Phi}_i}{N}, \widehat{\Phi}_{MG} = \frac{\sum_{i=1}^{N} \tilde{\beta}_i}{N}, \tilde{\lambda}_{jMG} = \frac{\sum_{i=1}^{N} \tilde{\lambda}_{ij}}{N}, \quad j = 1,\ldots,p\text{-}1$$

$$\hat{\delta}_{jMG} = \frac{\sum_{i=1}^{N} \tilde{\delta}_{ij}}{N}, j=0,\ldots,q\text{-}1, \ \hat{\theta}_{MG} = \frac{1}{N}\sum_{i-1}^{N} -({}^{\hat{\beta}_i}\!/_{\widehat{\Phi}_i}) \tag{8}$$

Pooled estimators are consistent and efficient assuming homogeneous long-run slopes. The homogeneity hypothesis can be tested empirically in all specifications and therefore not to be subjectively assumed. The existence of heterogeneity amongst the coefficients is examined by a Hausman-type test (Hausman, 1978).

The Dynamic Fixed effects estimator however is the most restrictive of all three estimators as it imposes long-run and short-run homogeneity of the slope parameters across countries, leaving only the intercepts to vary between cross-sections. In the DFE model $(N - 1)$ $(2k + 2)$ restrictions are imposed on the unrestricted model in equation (4), which amounts to k long-run coefficients, k short-run coefficients, the convergence coefficient and the common variance. However, Pesaran et al. (1999) points out that estimates of the mean values of the parameters in dynamic panel data models produced by the DFE estimators can be inconsistent and potentially misleading unless the slope coefficients are exactly identical.

3.2.5 Cross-Section Dependence in Panel Data

Panel-data models have a likelihood of exhibiting considerable dependence in the cross-sectional errors, which is likely to arise due to the existence of homogenous shocks and unobserved components which ultimately become incorporated in the residual term. In an estimation, the effect of cross-sectional dependence is as a result of a variety of factors, such as the degree of the correlations across cross sections and the type of cross-sectional dependence itself (De Hoyos and Sarafidis, 2006).

Dynamic panel estimators are prone to more severe problems arising from cross-sectional dependence than static panel estimators. When substantial cross-sectional dependence in the data is overlooked, it can lead to significant efficiency losses of such a magnitude that pooled estimators may provide little gain over the single-equation ordinary least squares (Phillips and Sul,2003).

In order to execute tests for cross-sectional dependence when $T< N$ three different testing procedures can be employed (Pesaran, 2004; Friedman, 1937; Frees ,1995) These tests are all valid when $T<N$. However, when $T>N$ which matches the panel data structure of the present study, the LM test, developed by Breusch and Pagan (1980) would be the preferable test procedure.

Assuming the standard panel-data model;

$$y_{it} = \alpha_i + \beta^{'}\mathbf{x}_{it} + u_{it} , \qquad i = 1,...,N \text{ and } t= 1,...,T \tag{9}$$

Where $\mathbf{x}_{it}$ denotes a $K\times 1$ vector of exogenous variables, β represents a $K\times 1$ vector of parameters under investigation, and α_i denotes individual nuisance parameters which are invariant to time. The null hypothesis assumes u_{it} to follow a gaussian white noise process within periods and between cross-sectional units. Under the alternative however, u_{it} may have some correlations between cross sections without violating the no serial correlation assumption.

Breusch-Pagan (1980) Lagrange Multiplier Test

Breusch and Pagan (1980) proposed an LM statistic to ascertain the null of zero cross-equation error correlations within a SURE framework which is defined as,

$$CD_{lm}=T\sum_{i=1}^{N-1}\sum_{j=i+1}^{N}\hat{\rho}_{ij}^{2} \tag{10}$$

In (10) above, $\hat{\rho}_{ij}$ denotes the sample estimate of the pair-wise residual Pearson correlation coefficient.

$$\hat{\rho}_{ij} = \hat{\rho}_{ji} = \frac{\sum_{t=1}^{N-1} e_{it}e_{jt}}{\left(\sum_{t=1}^{T} e_{it}^{2}\right)^{1/2}\left(\sum_{t=1}^{T} e_{jt}^{2}\right)^{1/2}} \tag{11}$$

where e_{it} is the OLS estimate of u_{it} in (11). LM's asymptotic distribution under the null hypothesis is chi-squared with N $(\frac{N-1}{2})$ degrees of freedom, as $T\rightarrow\infty$ with N fixed.

Pesaran's CD Test of cross-sectional dependence

Pesaran (2004) developed the following test,

$$CD = \sqrt{\frac{2T}{N(N-1)}}\left(\sum_{i=1}^{N-1}\sum_{j=i+1}^{N}\hat{\rho}_{ij}\right) \tag{12}$$

Under the null hypothesis of no cross-sectional dependence $CD \xrightarrow{d} N(0,1)$ for $N\rightarrow\infty$ and large T.

Pesaran, Ullah and Yamagata (2008) Bias adjusted LM test for cross-sectional independence

Pesaran, Ullah and Yamagata (2008) employing finite sample approximations rescale and re-cent the CD_{lm} test. The new LM test, denoted as PUY's LM test, is specified thus;

$$PUY's\ LM = \sqrt{\frac{2}{n(n-1)}} \sum_{i=1}^{n-1} \sum_{j=i+1}^{n} \frac{(T-K)\breve{\rho}_{ij}^2 - \mu_{Tij}}{\sigma_{Tij}} \quad (13)$$

where, $\mu_{Tij} = \frac{1}{T-K} tr\left(M_i M_j\right)$

is the exact mean of $(T-K)\breve{\rho}_{ij}^2$ and,

$\sigma_{ij}^2 = [tr\left(M_i M_j\right)]^2 a_{1T} + 2tr[(M_i M_j)^2] a_{2T}$

represents its exact variance. Here ,

$a_{1T} = a_{2T} - \frac{1}{(T-K)^2}$, a_{2T} =3$[\frac{(T-K-8)(T-K+2)+24}{(T-K+2)(T-K-2)(T-K-4)}]^2$

$M_i = I - X_i(\acute{X}_i X_i)^{-1}\acute{X}_i$

where T observations on the k regressors for the *i*-th individual regression is contained in $X_i = (x_{i1}, \dots, x_{iT})'$. PUY's LM is asymptotically distributed as *N* (0,1); under the null, with $T \to \infty$ first, and then $n \to \infty$.

3.3 Estimation Results

Results are provided in Table 1. The LLC and Breitung, Fisher type ADF, PP and IPS (1997) panel unit root tests validate that all data series are nonstationary I (1) processes. Estimation results are presented for the panel data unit root tests, panel cointegration tests and the MG, PMG and DFE estimations.

Tests were undertaken for the period 1997-2014 on an annual basis for the selected 8 Asian countries. The numbers in parentheses denote probability values. The Fisher-type test follows a chi-squared distribution under the null hypothesis, whereas asymptotic normality is assumed for the others. Table 2 shows the Pedroni (1999) cointegration test, following Pedroni (1999), the first two statistics which are denoted panel cointegration statistics are grounded on the within approach and assume heterogeneity while the last two are denoted cointegration statistics of the group panel framework, which follow the between approach and assume heterogeneity. All the statistics of the different tests follow normal distributions and their values (calculated) are comparable with the Pedroni (1999) critical values. The results of the within and between panel cointegration tests in summary shows that the null hypothesis of no cointegration are rejected in panel PP- statistic, Panel ADF- statistic, group PP-statistic and Group ADF-statistic. Consequently, the assumption of a long-run equilibrium relationship between employment rate, investment, wage, production and R&D expenditure is thus validated.

Table 3 and 4 presents the estimates of the long-run and the short-run coefficients, the adjustment coefficient, cross sectional independency and the Hausman test statistics. The Hausman statistic indicated a value of (52.9 (0.000)) and (14.61(0.0122)) for MG/PMG in the monotonic and non-monotonic models respectively. As a result, the homogeneity of the long-run coefficients across industries is not supported by the model making the MG model the more consistent and preferred model. The Hausman statistic, (0.56(0.9677)) and (4.60(0.4670)) for MG/DFE in the monotonic and non-monotonic models respectively validates the non-existence of the simultaneous equation bias emanating from the endogeneity between the error term and the lagged dependent variable as discussed in Baltagi et al (2000), as such the homogeneity of the short and long-run parameters across countries in the linear model is supported making the DFE model the more efficient and preferred model.

The long run coefficients for the linear model indicates significance for the r&d variable only in the MG model and weak significance in the PMG model, the r&d coefficient signs were different for both models with the MG model having a negative r&d coefficient and the PMG model having a positive one. The DFE model however had largely insignificant long-run coefficients even though there was weak evidence for short-run adjustment. Short-run adjustment was however evident in the MG model and largely insignificant in the PMG model. Due to the conflicting nature of the results across models bordering on differences in coefficient signs it becomes rather pertinent to impose non-linear effects within the model space by introducing a quadratic version of the r&d variable. Incorporating non-linearity also arises from the fact that in the literature, the coefficient sign of the r&d variable is different across industry specific firms and our study makes no distinction between r&d variables across industry specific firms given the aggregate nature of our r&d variable. The estimation results are outlined in table 4.

In order to mitigate the effects of potential multicollinearity, the r&d and r&d^2 variables were standardized prior to estimation. An interesting scenario opens up in table 4 with the non-linear hypothesis being supported in the PMG and DFE models but not attaining significance in the MG model. The inability of the MG model in attaining significance maybe owing to the loss of degrees of freedom within each cross-section. However, the efficiency gains obtained by pooling in the PMG and DFE models maybe the reason significant r&d coefficients were obtained in both models as more observations were used for estimation in those models unlike in the MG model that estimates separate regressions for each cross-section. In the long run, both the PMG and DFE models indicate a local extremum of 3.4 and 0.4. The minimum point for the PMG model was outside the data range but was well within the data range for the DFE model. The results are not unexpected considering that the Hausman test for MG/PMG shows the model to be largely inconsistent and inefficient while that of the MG/DFE showed the DFE model to be more efficient and consistent. The wage and GDP variables for the PMG model followed a priori expectations in both the linear and non-linear models as the wage variable was negative and significant indicating a negative labour cost effect, while the GDP variable indicated a significant and positive income effect. Capital formation however was largely insignificant for all models in both linear and non-linear specifications. The linear and non-linear MG model had mostly insignificant long- run coefficients except for the GDP variable that indicated a significant and positive income effect for the linear MG model. The same is also applicable for the DFE model which obtained significant r&d coefficients in the non-linear model to support the non-linear U shape relationship between r&d expenditure and employment. The results indicate that lower levels of r&d expenditure contracts employment but would induce employment within the 75th percentile of the r&d data range. The non-linear specification appeared to be a bit more robust than the linear specification due to the uniformity of r&d coefficient signs.

A positive short-run income effect was indicated for all models in the linear specification, although all the other variables were largely insignificant for all models in the linear short-run specification. The significance of GDP was however lost in the non-linear specification of the PMG model but was retained in the MG and DFE models. In addition, capital formation attained short-run significance in the non-linear specification for both the MG and DFE models but not in the PMG model. Capital formation indicated short-run negative effects for both models. This is indicative of a short-run labour saving capital effect as a result of process innovation, consistent with Piva and Vivarelli (2017)

The results of cross section dependency via the Breusch-Pagan test supports the null hypothesis non-rejection of cross-sectional independence.

Table 1: Tests for unit roots for the panel (1997-2014).

Variables	LLC t-test	IPS t-test	ADF-Fisher chi-square	PP-Fisher chi-square
Level				
L	-1.42473 *	-0.62992	21.7371	16.6128
	(0.0771)	(0.2644)	(0.1519)	(0.4111)
Lgi	-2.55533 ***	-0.08726	18.7351	4.97660
	(0.0053)	(0.4652)	(0.2826)	(0.9959)
r&d	1.43101	-0.05068	30.0459**	30.9846**
	(0.9238)	(0.4798)	(0.0178)	(0.0135)
Lw	-1.59941 *	-0.29570	17.3867	7.76886
	(0.0549)	(0.3837)	(0.3610)	(0.9555)
Lgdp	-3.03271***	0.64134	16.8581	5.11714
	(0.0012)	(0.7393)	(0.3948)	(0.9951)
First difference				
Δl	-2.99705***	-3.16064***	38.0463***	68.1796***
	(0.0014)	(0.0008)	(0.0015)	(0.0000)
Δlgdp	-6.46572 ***	-5.75195 ***	61.4597 ***	60.9943 ***
	(0.0000)	(0.0000)	(0.0000)	(0.0000)
ΔLgi	-2.44083 ***	-3.68032 ***	40.8018 ***	70.3974 ***
	(0.0073)	(0.0001)	(0.0006)	(0.0000)
Δr&d	-8.78731 ***	-7.98398 ***	84.9545 ***	230.807 ***
	(0.0000)	(0.0000)	(0.0000)	(0.0000)
ΔLw	-9.13670 ***	-7.54768 ***	78.9758***	78.3061 ***
	(0.0000)	(0.0000)	(0.0000)	(0.0000)

Note: "***", "**", "*", indicates the significance of the estimated parameters at the 1%,5% and 10% levels, respectively. L denote natural logarithms. Δ denotes the first difference of logarithm form of variables.

Table 2: Cointegration test for the 8 Asian countries (1997-2014) for total industry.

homogeneous			heterogeneous		
Test	Statistics	Prob.	Test	Statistics	Prob.
Pedroni residual cointegration					
Panel PP-Statistic	-3.019***	0.0013	Group PP-statistic	-8.86615***	0.000
Panel ADF-statistic	-3.465***	0.0003	Group ADF-Statistic	-5.06658***	0.000

Note: (***) denotes that the parameter estimates are significant at the 1% level.

Table 3: Panel estimation for 8 Asian countries (1997-2014), linear model.

Variables	MG	PMG	DFE
Long-run			
Lgi	-0.7677581	0.1892387	-7.903569
	(0.920)	(0.943)	(0.511)
Lgdp	21.14121**	16.25616***	20.0831
	(0.032)	(0.000)	(0.330)

Lw	5.295213	-0.9129518***	1.637665
	(0.202)	(0.001)	(0.502)
r&d	-3.202531***	1.514562*	3.150866
	(0.006)	(0.081)	(0.634)
Short-run			
adjustment coefficient	-0.5274966***	-0.2128445	-0.0513779*
	(0.008)	(0.137)	(0.078)
ΔLgi_{t-1}	-3.631987	-2.766026	-1.833454
	(0.170)	(0.177)	(0.160)
$\Delta Lgdp_{t-1}$	14.07998**	13.37557***	13.32976***
	(0.026)	(0.001)	(0.000)
ΔLw_{t-1}	0.3875292	0.3428751	-0.0230696
	(0.625)	(0.404)	(0.881)
$\Delta r\&d_{t-1}$	-0.3715431	-0.590884	0.5547804
	(0.818)	(0.636)	(0.376)
Hausman test (MG/PMG) 52.94*** (0.000) MG/DFE) 0.56(0.9677)			
Cross-sectional independence for quadratic specification			
LM	32.54		
	(0.2533)		
LM adj	0.0848		
	(0.9325)		
LM CD	-0.1845		
	(0.8536)		

Notes: "***", "**", "*", indicates significance at the 1%, 5% and 10% levels respectively. Δ denotes the first difference while l denotes the logarithmic form of the variables.

Table 4: Panel estimation for 8 Asian countries (1997-2014) non-linear model.

Variables	MG	PMG	DFE
Long-run			
Lgi	-13.191	0.82950	-1.5887
	(0.264)	(0.468)	(0.738)
Lgdp	5.0835	6.2482***	10.7515
	(0.251)	(0.000)	(0.170)
Lw	2.0667	-1.373***	0.86171
	(0.329)	(0.003)	(0.378)
r&d	-17.317	-14.829**	-27.299***
	(0.705)	(0.046)	(0.000)
$r\&d^2$	17.505	101.198**	19.954***
	(0.912)	(0.032)	(0.000)
Short-run			
adjustment coefficient	-0.6796815***	-0.1713419	-0.1229563***
	(0.000)	(0.102)	(0.001)
ΔLgi_{t-1}	-6.153211**	1.133161	-2.497583*
	(0.040)	(0.800)	(0.052)
$\Delta Lgdp_{t-1}$	19.41794***	8.491269	14.43215***
	(0.004)	(0.264)	(0.000)
ΔLw_{t-1}	0.5341985	-0.0317694	0.0359838
	(0.239)	(0.945)	(0.814)

$\Delta r\&d_{t-1}$	8.734818 (0.665)	9.768922 (0.296)	2.951876 (0.263)
$\Delta r\&d^2_{t-1}$	-105.0448 (0.665)	-15.37026 (0.619)	-0.9863409 (0.521)
Hausman test (MG/DFE) 4.60(0.4670)			(MG/PMG) 14.61**(0.0122)
Cross-sectional independence test for quadratic specification			
LM	39.47 (0.0737)		
LM adj	1.555 (0.1198)		
LM CD	-0.9059 (0.3650)		

Notes: "***", "**", "*", indicates significance at the 1%, 5% and 10% level., Δ denotes the first difference while l denotes the logarithmic form of the variables.

3.4 Discussions of Results

The uniformity of the robustness of the non-linear estimation results allude to the suggestion that innovation may indeed have a non-linear relationship with employment. This result may actually be dependent on a combination of scale and technique effects. At lower levels of product innovation, jobs get destroyed at a faster rate than they are created due to the fact that the supply side dimension is not as developed as the demand side. The products are being created at a lower scale than they are being used so the job destroying effect tends to outpace the job creating effect. Another reason for this may be due to the fact that once these new products are developed there usually are not enough workers with the necessary skill set required to produce the new products. Another important critical element in this argument may be that the skillset required to produce the new products may be higher than the skillset required for the jobs these products were intended to replace in the first place. Also, the new products may be less labour intensive than the jobs they were designed to replace. This entails a longer time period required to get the supply side to be at par with the demand side. Once product innovation attains a higher rate, the supply side comes up to scale. This would naturally necessitate the employment of labour to produce the new products. Also, job creation due to product innovation can be both direct and indirect. Direct job creation would naturally be due to the need for labour in the production of the new products. Indirect job creation would entail the development of forward and backward linkages. Forward linkages are developed due to products that may originate from the utilization of the original product innovations. A very good instance would be the software market which is a forward linkage to the personal computer market. Backward linkages on the other hand may develop due to an increased demand for the raw materials which are utilized in the production of the original product innovations. Raw materials in this instance also includes intermediate and finished goods. The development of these linkages can also induce job creation. It is then safe to deduce from the above arguments that— as the intensity of R&D in the economy increases so also would forward and backward product linkages which could potentially induce employment increase.

3.5 Conclusion and Policy Implications

The study employs panel cointegration methods of Pedroni (1999) and panel estimation methods of Pesaran et al (1999) and Pesaran and Smith (1995) in order to analyze the innovation and employment nexus in a panel of 8 Asian countries within the period 1997-2014. By utilizing linear and non-linear specifications a non-linear innovation-employment relationship was uncovered with a local maximum point at about the 75th percentile of the r&d data range. The underlying ramifications of the result entails that innovation contracts employment at close to median values and that at higher values, r&d tends to improve employment. The contractionary effect of lower values of r&d maybe due to the increased skills requirement in jobs which are created due to product innovation or underlying process innovation embedded in r&d expenditures as what may be product innovation in one sector may turn out to be process innovation in another (Dosi, 1984). Policy implications would include incentivizing the creation of products which have the highest potential linkages in order to induce further employment from product innovation.

Chapter 4

INNOVATION AND INTERSECTORAL EMPLOYMENT IN EUROPEAN COUNTRIES

4.1 Introduction

This paper estimates the impact of research and development (R&D) on employment at the industrial level. The path of this effect remains quite vague in the theoretical literature and thus needs to be verified empirically. The dataset aids us to distinguish the impact of R&D on employment over the sectoral dimension for EU-12 countries for the 2000-2009 period for the manufacturing sectors (High-Tech and Low-Tech) and the 2000-2015 period for industry and service sector based on data availability. Our method allows us to ascertain the long-run cointegrating coefficients. Analyzing the possible employment inducing impact of technological innovation is an old and contentious topic. The dispersion of the "new economy" grounded on information and communication (ICT) technologies for about a generation now has triggered a recurrence of the age-old discussion on the potential employment (creating) effects of innovation. At the advent of the industrial age, the fear of being dismissed loomed large in the working-class circles due to technological change (Ricardo, 1951) whilst the academia were deeply confident about the market compensation mechanism to correct for workers' dismissal. Labor-saving innovation which may stimulate technological unemployment has always been viewed with suspicion by the labor force in periods experiencing radical technological change[1]. From the perspective of macroeconomics (aggregates) , the direct labor saving effect of process innovation has to be linked to the theory that highlights the existence of indirect effects (product innovation and price and income effects) which could offset the drop in employment, as a result of the incorporation of process innovation in the new machineries. About two centuries ago, a theory was put forward by classical economists which was later styled the "compensation theory "(Marx, 1961). Both Ricardo and Marx's inferences regarding the effect of machines on labourers. Mainly, while Ricardo and Marx explain the impact of mechanisation in any production line on labour demand, they concurringly mention the fact that when the fixed capital or floating capital rates change in the capital of any production line, the labourers will be dismissed from their jobs at that production line (Ricardo,1951; Marx, 1961*)*. However, the theory which suggests that the capital which is freed after the labourers are dismissed from their jobs, will later employ the same labourers, and somehow "compensate" the dismissal is being supported by classical economists such as (Stuart Mill in; Marx,1961; Senior,1836), but there is no clear explanation about the extent of this compensation. (Marx,1961) also gives support for the "compensation theory", but makes a case for the resultant negative effect on worker's welfare.

The components of this compensation mechanism are primarily price and income effects. Process innovation puts a downward pressure in prices by stimulating a decline in production cost. In markets with competition, this increases aggregate demand and an upsurge in productivity levels and employment. With regards to the income effects,

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1 . The mythical activities of Ned Ludd in industrial areas and Capitain Swing in the countryside on industrial facilities can lay testament to this fact (see Hobsbwm, 1968; Hobsbawm and rude, 1969).

in a world where competitive convergence is not contemporaneous, it would be observed that during the delay between cost reductions arising from innovation (process) and the subsequent price reductions, additional revenues and/ or additional earnings accrued by the innovators and their workers may be reinvested and thus may spur additional employment and/or increase aggregate demand due to higher consumption which may also lead to the creation of more jobs (Stoneman, 1983), this can replenish the previous job losses owing to process innovation (Boyer, 1988). Also, technological change does not always translate to process innovation. It can also entail the creation of completely new economic areas where new jobs can be created. In contemporary debate, various studies (Freeman, 1982; Freeman and Soete, 1987; Freeman and Soete 1994; Vivarelli, Evangelista and Pianta, 1996; Vivarelli and Pianta, 2000) all agree that product innovations generate employment since they lead to the creation of either entirely novel products or main modification of already existing ones. Since the employment outcome of innovation is not clearly defined by economic theory, there arises a great necessity for an empirical investigation to calibrate the resulting employment effect of technological change.

Following this framework this chapter of the thesis aims to empirically investigate the potential job creating impact of business R&D expenditures at the aggregate industrial level in the long-run[2]. This study contributes to the existing literature in three ways: Firstly, the macroeconomic investigation is based on panel dataset, able to overcome

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2. Note that the innovation variable used in this study (R&D) is a preferred proxy of product innovation rather than of process innovation. While process innovation is mainly incorporated in the new modifications of fixed capital, R&D is mainly aligned towards the promotion of prototypes, the introduction of entirely new products, or the radical differentiation of existing products (see Conte and Vivarelli (2005), Bogliacino & Piva & Vivarelli (2011), Bogliacino & Vivarelli (2012).

the limitations of previous empirical studies, mainly based on either Firm-level cross-section/panel analysis or single country data. Secondly, the proxy variable for technology is measurable and continuous, while a great number of past studies have captured technological change via indirect proxies or indicator variables (for instance, the incidence of product or process innovation). Thirdly, the dataset controls for the capturing of the sectoral employment inducing impact of aggregate R&D, with the option of focusing on low-tech, high tech manufacturing and services sectors. Therefore, we are able to separate the advent of job-creating effects across the different segments of the economy. To the best of our knowledge, sectoral comparisons have been carried out in very few past studies (Bogliacino and Vivarelli, 2012; Mastrostefano and Pianta, 2009; Van roy. et al 2015; Aldieri and Vinci.P.C, 2017). Fourthly, the panel fully modified ordinary least squared (FMOLS) method used in our study allows us to estimate long-run coefficients of the cointegrating relationship within EU-12 countries. To the best of our knowledge this will be the first study to take into consideration non-stationarity in the relationship between R&D and employment.

4.2 Data and Methodology

Data for yearly, wage, R&D expenditure, investment and employment are gathered from OECD STAN and OECD ANBARD, they were chosen based on data availability for EU-12 countries for the period of 10 years (2000-2009) for manufacturing sectors (both high-tech and low-tech) and for services and industry period of data is 15 years (2000-2015). All independent variables collected in national currencies and deflated by Consumer Price Index (CPI). Preceding the empirical analysis, the dataset are logarithmically transformed (natural logarithms). The selected EU-12 countries for manufacturing sectors are: Austria, Belgium, Czech Republic, France, Germany, Greece, Hungary, Italy, Netherlands, Slovenia, Spain and Norway and the countries selected for service sector and industry are: Belgium, Czech Republic, France, Germany, Italy, Slovenia, Spain, Norway, Finland, Poland, Portugal and United Kingdom. Because of missing data from some countries, we obtain different categories of data from different countries.

4.2.1 The Employment Demand Equation

The employment demand equation in its implicit form can be modelled as:

lemp =f (lgi, lw, lR&D) (14)

Where l stands for employment, gi stands for gross fixed capital, w stands for wage and R&D stand for Business research and development expenditure. All variables are collated in local currency values and transformed into real values before being logarithmically transformed.

The stochastic version of the model by the inclusion of innovation (see for similar approach, Lanchenmaer and Rottmann, 2011, Bogliacino et al., 2011, Vincent et al., 2015) for a panel of country (i), over time (t) is:

$$l\,emp_{it} = \beta_0 + \beta_1\, l\, w_{it} + \beta_2 lr\&d_{it} + \beta_3\, lgi_{it} + v_{i,t} \qquad (15)$$

where v is the usual error term.

4.2.2 Panel Unit root Test

The stationarity of each variable mentioned in Eq.14 was verified through unit root tests: The Fisher-ADF tests of Maddala and Wu (1999) and Choi (2001), LLC test of Levin et al (2002) and IPS test of IM et al (2003) were all employed for the tests. It has been emphasized that single country tests may exhibit low power (Harris and Sollis, 2003). Two approaches for analysing the non-stationarity tests in levels and first differences are examined. A typical equation for the panel unit root that is represented by Levin et al. (Dicky and Fuller, 1979) is denoted as:

$$\Delta y_{it} = \alpha_i + \beta_i y_{it-1} + \sum_{j=1}^{pi} \rho_i \Delta y_{it-j} + e_{it} \tag{16}$$

where ρ is chosen in such way that the residuals are uncorrelated across time, Δy_{it} indicates y_{it} in first differences for country denoted by i, in time periods denoted by t=1,…,T. Since the LLC technique follows the homogeneous panel postulation, β_i is equal across all countries. The null hypothesis is that H_0: $\beta_i = 0$ against the alternative is thatH_1 $\beta_i > 0$ for some i, which assumes mean reversion or stationarity in all the series.

4.2.3 Panel Cointegration Test

We employed cointegration technique developed by Pedroni (1999) for panel data to determine possible long-run relationships amongst employment rate and the selected independent variables. In contemporary studies, analysis of cointegration and panel estimation methods, including studies by Larrsson et al. (2001), Kao (1999), Harris and Sollis (2003) have been accorded great attention. In their studies, handling short-run dynamics and with regards to heterogeneity and cross-sectional heterogeneity remains a very contentious issue, for the heterogeneity issue appears to predominate panel data analysis. The analysis for cointegration and panel estimation technique of Pedroni (1995, 1999, 2000) is employed to circumvent problems arising from heterogeneity.

Additionally, the ADF-type test which is parametric has been shown to be super consistent when processes with autoregressive (AR) errors are being modelled, reason being that the AR terms are more precisely captured. Accordingly, utilizing a different array of testing techniques can be useful if the underlying data generating process statistics are not known. The present study will employ cointegration tests using both homogenous and heterogenous procedures.

4.2.4 The Between-Group Panel FMOLS Estimator

To estimate the cointegrating vector, we will use the between-group FMOLS (panel cointegration estimator). FMOLS is a very frequently employed panel data estimation procedure. Its approach to estimating optimal cointegrating regression coefficients is non-parametric (Phillips & Hansen, 1990), due to the presence of cointegrating relationships it is modified to make adjustments for serial correlation and endogeneity (Phillips, 1995). Pedroni (2000,2001) puts forward two procedures for the application of this fully modified method to panel cointegration regression, the pooled (or within-group) panel FMOLS estimator and the group-mean (between-group) FMOLS estimator. Also, Pedroni maintains that between-group estimators perform better than the within-group estimators since it controls for a more flexible alternative hypothesis and is more invariant to small sample size distortion than the within-group estimator. The group-mean panel FMOLS estimator for Eq. (14) can be outlined as:

$$\beta^*{}_{fmols} = \frac{1}{N}\sum\left[\frac{\sum_{t=1}^{T}(logr\&d-\overline{logr\&d}\,)logl_{it}-T\gamma_{it}}{\sum_{t=1}^{T}(logr\&d-\overline{logr\&d})\hat{}2}\right] \tag{17}$$

4.3 Results

4.3.1 Unit Root Test Results

Two models have been employed to determine the presence of non0stationarity in the panel. The initial model is augmented with a deterministic trend as well as an intercept, while the later one controls for just intercept without any trend terms. Test outcomes are outlined in Tables 5-8, LLC and Breitung t-tests, Lee and Chiu (2011), the IPS t-test, and Fisher type ADF and PP chi-square tests have all been employed to ascertain the presence of stationarity in the panel. These testing techniques consider non mean reversion as the null hypothesis compared with the alternative of mean reversion

(Harris and Sollis, 2003). For all variables, the IMP t-test and the ADF tests reject the null hypothesis of non-stationarity at 1%, 5% and 10% significant level in the level form in industry, service and manufacturing sectors (low-tech and high-tech). However, the results were mixed for the LLC t-test and the PP-Fisher Chi-square test. by taking the difference, the null hypothesis is rejected at 1%,5%, 10% significance

Table 5: Panel unit root tests for panel data. Whole sample (industry).

Variables	LLC t-test	IMS t-test	ADF-Fisher chi-square	PP-Fisher chi-square
Level				
Ll	-1.74986 **	-0.67909	27.1062	34.4517 ***
	(0.0401)	(0.2485)	(0.2995)	(0.0770)
Lgi	-1.65274 **	-0.02614	21.9719	12.5434
	(0.0492)	(0.4896)	(0.5809)	(0.9731)
LR&D	-0.58749	1.34134	20.7507	11.5386
	(0.2784)	(0.9101)	(0.6534)	(0.9846)
Lw	-3.54540 *	-5.55536	28.3945	75.6135 *
	(0.0002)	(0.2893)	(0.2438)	(0.0000)
First difference				
ΔLl	-6.56015 *	-4.1509 *	57.2680 *	54.4696 *
	(0.0000)	(0.0000)	(0.0002)	(0.0004)
ΔLgi	-5.56002 *	-3.34693 *	47.9948 *	41.7112 *
	(0.0000)	(0.0004)	(0.0025)	(0.0139)
ΔLR&D	-8.15892 *	-4.48504 *	70.6249 *	93.1016 *
	(0.0000)	(0.0000)	(0.0000)	(0.0000)
ΔLw	-4.92711 *	-2.64562 *	31.3117	35.7787 **
	(0.0000)	(0.0041)	(0.1450)	(0.0577)

Note: (*), (**), (***) denotes test significance at the 1%, 5% and 10% levels, respectively. Δ denotes the first difference. L denotes the logarithmic form of the variables.

level. All the variables were tested with and without trend terms, as well as in level and first difference. Our test results mostly support the existence of unit roots at levels and its absence at first difference. In conclusion *lemp, lgi, lR&D* and *lw* all follow an I(1) process. Unit root test results are presented in Table 5-8. The results for the estimation are collated from unit root tests (panel) for the periods 2000-2015 at yearly periods for the EU-12 countries under investigation, having 132 observations in manufacturing sectors and 180 observations in industry and services sector. Values in parentheses represent probabilities.

Table 6: Panel unit root tests. Service sector.

Variables	LLC t-test	IMS t-test	ADF-Fisher chi-square	PP-Fisher chi-square
Level				
Ll	-1.74986**	-0.6909	27.1062	34.4517 ***
	(0.0401)	(0.2485)	(0.2995)	(0.0770)
Lgi	-2.04195 **	-0.34693	25.2594	12.2684
	(0.0206)	(0.3643)	(0.3918)	(0.9767)
LR&D	-0.95263	2.43770	8.48800	9.51171
	(0.1704)	(0.9926)	(0.9985)	(0.9963)
Lw	-14.19755*	-0.75113	32.0327	55.3109 *
	(0.0000)	(0.2263)	(0.1262)	(0.0003)
First difference				
ΔLl	-6.56015 *	-4.15059 *	57.2680 *	54.4696 *
	(0.0000)	(0.0000)	(0.0002)	(0.0004)
ΔLgi	-4.64372 *	-3.03618 *	45.2119 *	39.0522 **
	(0.0000)	(0.0012)	(0.0055)	(0.0270)
ΔlR&D	-7.50190 *	-6.05133 *	78.8446 *	84.9977 *
	(0.0000)	(0.0000)	(0.0000)	(0.0000)
ΔLw	-2.82531 *	-1.07263	34.2968 ***	29.4668
	(0.0024)	(0.1417)	(0.0795)	(0.2031)

Note: (*), (**), (***) denotes tests significance at the 1%, 5% and 10% levels respectively. Δ denotes the first difference. L denotes the logarithmic form of the variables

Table 7: Panel unit root tests (2000-2009) for Low-Tech manufacturing sector.

Variables	LLC t-test	IMS t-test	ADF-Fisher	PP-Fisher
Level				
Ll	-3.89250*	-0.59252	26.4877	22.2589
	(0.0000)	(0.2768)	(0.3289)	(0.5638)
Lgi	-1.68058 **	-0.12253	23.2210	57.0459 *
	(0.0464)	(0.4512)	(0.5068)	(0.0002)
LR&D	0.79207	0.37942	20.9511	41.7354*
	(0.7858)	(0.6478)	(0.6416)	(0.0137)
Lw	-0.66079	2.16738	12.3093	10.7135
	(0.2544)	(0.9849)	(0.9762)	(0.9909)
First difference				
ΔLl	(-3.69630) *	(0.21041)	(27.6200)	(37.9647) **
	0.0001	0.5833	0.2764	0.0350
ΔLgi	(-6.30284) *	(-3.28074) *	(55.9385) *	(89.9953) *

	0.0000	0.0005	0.00002	0.0000
ΔlR&D	(-13.8800) *	(-6.54815) *	(91.0986) *	(110.051) *
	0.0000	0.0000	0.0000	0.0000
ΔLw	(-8.10192) *	(-2.64562) *	(47.9863) *	(68.8232) *
	0.0000	0.0041	0.0025	0.0000

Note: (*), (**), (***) denotes tests significance at the 1%, 5% and 10% levels respectively. Δ denotes the first difference. L denotes the logarithmic form of the variables

Table 8: Panel unit root tests (2000-2009) for High-Tech manufacturing sector.

Variables	LLC t-test	IMS t-test	ADF-Fisher	PP-Fisher
Level				
Ll	-0.93738	0.11190	(20.839	15.6819
	(0.1743)	(0.5445)	(0.5307)	(0.8314)
Lgi	-3.8454 *	-0.74831	31.5020	74.5370 *
	(0.0001)	(0.2271)	(0.1398)	(0.0000)
LR&D	-1.06558	3.19728	6.93391	12.3359
	(0.1433)	(0.9993)	(0.9997)	(0.9759)
Lw	-1.15162	1.30855	20.0400	12.3359
	(0.1247)	(0.9047)	(0.6944)	(0.9759)
First difference				
ΔLl	-6.64438 *	-2.91093 *	47.5248 *	52.7441
	(0.0000)	(0.0018)	(0.0013)	(0.0002)
ΔLgi	-10.3052 *	-3.13427 *	53.4746 ***	112.719 *
	(0.0000)	(0.0000)	(0.0005)	(0.0000)
ΔlR&D	-8.87490 *	-4.66738 *	68.4363 *	81.9347 *
	(0.0000)	(0.0000)	(0.0000)	(0.0000)
ΔLw	-4.72557 *	-001192	24.4630 *	38.4467 **
	(0.0000)	(0.4952)	(0.4354)	(0.0312)

Note: (*), (**), (***) denotes tests significance at the 1%, 5% and 10% levels respectively. Δ denotes the first difference. L denotes the logarithmic form of the variables

4.3.2 Panel Cointegration Results

As indicated earlier, the following process examines the long-run relationship amongst unemployment rate, investment, production, R&D expenditure and wage in industry and the three sectors (services, high-tech and low-tech manufacturing) The results of the panel cointegration tests are outlined in Table 9- 12. For all the variety of tests, it is important to know that all of the statistics do not follow standard distributions and thus, the obtained values (calculated) are compared with critical values obtained from

Pedroni (1999). To summarize, the result of homogeneous panel cointegration indicates that the no cointegration null hypothesis is rejected in panel PP- statistic and Panel ADF- statistic. Rejection in the heterogeneous panel cointegration is supported for group statistic (PP and ADF) for both total industry and the sectors (service, high and low-tech manufacturing). Consequently, there exists a long-run equilibrium relationship between unemployment rate and investment, wage and R&D expenditure in both industry and sectors (services, low-tech and high-tech manufacturing). The long-run elasticity for unemployment rate and independent variables are obtained through FMOLS estimation which follows subsequently in the study.

Table 9: Cointegration test for EU-12 countries (2000-2015) for industry sector.

Within dimension (homogeneous)			Between dimensions (heterogeneous)		
Statistic	value	Prob.	Statistic	value	Prob.
Panel v-Statistic	-0.36366	0.6419	Group rho-Statistic	1.754468	0.960
Panel rho-Statistic	0.355599	0.6389	Group PP-statistic	-4.8157*	0.000
Panel PP-Statistic	-3.4261*	0.0003	Group ADF- Stat	-3.8429*	0.000
Panel ADF-statistic	-3.3386*	0.0004			

Note: (*) indicates that the estimated parameters are significant at the 1% level.

Table 10: Cointegration test for EU-12 countries (2000-2015) for service sector.

Within dimension (homogeneous)			Between dimensions (heterogeneous)		
Statistic	value	Prob.	Statistic	value	Prob.
Panel v	0.83812	0.2010	Group rho	2.00980	0.9778
Panel rho	0.63607	0.7376	Group PP	-5.2547*	0.0000
Panel PP	-3.5125*	0.0002	Group ADF	-5.6951*	0.0000
Panel ADF	-4.5167*	0.0000			

Note: (*) indicates that the estimated parameters are significant at the 1% level

Table 11: Cointegration test for EU-12 countries (2000-2009) low-tech manufacturing sector.

Within dimension (homogeneous)			Between dimensions (heterogeneous)		
Statistic	value	Prob.	Statistic	value	Prob.
Panel v-Statistic	-1.21826	0.8884	Group rho-Statistic	4.220313	1.0000
Panel rho-Statistic	3.390585	0.9997	Group PP-statistic	-5.8470*	0.0000
Panel PP-Statistic	-1.8082**	0.0353	Group ADF-Statistic	-2.5349*	0.0056
Panel ADF-statistic	-1.3505***	0.0884			

Note: (*), (**), (***) denotes tests significance at the 1%, 5% and 10% levels respectively. Δ denotes the first difference. L denotes the logarithmic form of the variables

Table 12: Cointegration test for EU-12 countries (2000-2009) High-Tech manufacturing sector.

Within dimension (homogeneous)			Between dimensions (heterogeneous)		
Statistic	value	Prob.	Statistic	value	Prob.
Panel v-Statistic	-1.2939	0.9022	Group rho-Statistic	3.6574	0.9999
Panel rho-Statistic	2.2501	0.9878	Group PP-statistic	-1.847**	0.0324
Panel PP-Statistic	-1.5886**	0.0561	Group ADF-Statistic	-1.0976	0.1362
Panel ADF-statistic	-1.3413***	0.0899			

Note: (**), (***)denotes tests significance at the 5% and 10% levels respectively.

4.3.3 FMOLS Estimation Results

Tables 13-16, presents the panel fully modified least squares (FMOLS) model that is based on panel yearly data for the period 1999-2009 with 132 observations for manufacturing sectors and for the period (2000-2015) with 180 observations for industry and service sector. In tables 13-16 Pedroni (1999) between and within procedures for estimation have been used for the FMOLS panel result. As we can see from the total sample in Table 13, the long-run elasticity (0.031) is statistically significant at the 1% and 5% levels. Panel long-run elasticity for lgi, lR&D and lw variables are shown to exhibit inelasticity and the signs for the lgi and lR&D

coefficients are positive and for lw is negative, which is consistent with apriori expectations except for wage and capital in the low-tech manufacturing sector that are insignificant. Shifting our focus to the main variable of interest (R&D) and employment. Generally, the evidence not only shows support for a labour-friendly role of R&D expenditure but the effect is small. If an industry doubles the size of its R&D expenditure, the projected increase in employment within that industry would be 0.030586% and 0.030468% (for homogenous and heterogenous correspondingly). To probe the association between innovation pressures and intersectoral employment levels, the specification was employed for various subsamples. Tables 9-12 are results of industry, services, low and high-tech manufacturing sectors. From observations, the econometric results of demand for labour (production, investment and wage) are consistent across all the tables. Therefore, our comments will emphasise only the R&D expenditure coefficient. Indeed, the employment impact is highly confirmed in the case of the manufacturing industry (with magnitude of 0.030586 at a 99% level of confidence for homogeneous and with 0.030468 magnitude at 99% confidence level for heterogeneous), while it is consistently confirmed in the case of services (with a magnitude of 0.039741 at 99% for homogenous and 0.121052 at 95% for heterogeneous). once we isolated the low-tech manufacturing sectors from the high-tech manufacturing sectors. we find insignificant coefficients (-0.039297 for homogeneous and -0.042001 for heterogeneous) (with the losses for gross investments) for low-tech manufacturing sector and for high-tech manufacturing sector we find insignificant coefficients–although still positive in homogeneous (with a magnitude of 0.027917) but highly significant in the heterogeneous panel (with a magnitude of 0.756212 at a 99% level of confidence).This evidence lends credence to the opinion that innovation is more employment friendly in the services and high-tech manufacturing sectors while not significant in the low-tech manufacturing sectors. This outcome can be due to the fact that both sectors (high-tech production and services) are categorized by a foremost role of production innovation which engenders more efficient “compensation mechanisms” driven via growing demand while the more traditional low-tech manufacturing sectors are in contrast characterized via a preponderance of process innovation and demand reductions.

Table 13: Panel (FMOLS) estimation EU-12 countries whole sample (industry)

Variables	FMOLS	t-statistic	Prob.
Heterogeneous			
Lgi	0.189132*	(22.77182)	0.0000
LR&D	0.030468*	(6.678182)	0.0000
Lw	0.024820	(1.599623)	0.1118
Homogeneous			
Ly	0.182987*	(7.229927)	0.0000
LR&D	0.030586**	(2.199969)	0.0294
Lw	0.034423	(0.728009)	0.4678

Note: (*), (**) denotes tests significance at the 1% and 5% levels respectively.

Table 14: Panel (FMOLS) estimation EU-12 countries (service sector)

Variables	FMOLS	t-statistic	Prob.
Heterogeneous			
Lgi	0.179340*	(4.334187)	0.0000
LR&D	0.121052**	(2.568074)	0.0112
Lw	-0.119468*	(-3.336039)	0.0058
Homogeneous			
Ly	0.207032*	(8.552728)	0.0000
LR&D	0.039741*	(4.215763)	0.0000
Lw	-0.14594*	(-2.901896)	0.0043

Note: (*) denotes tests significance at the 1% level

Table 15: Panel (FMOLS) estimation EU-12 countries (low-tech manufacturing sector).

Variables	FMOLS	t-statistic	Prob.
Heterogeneous			
Lgi	-0.008161	(0.597367)	0.5517

LR&D	-0.042001*	-7.273555	0.0000
Lw	-0.115982*	(-4.485044)	0.0000
Homogeneous			
Ly	-0.027716	(-0.946050)	0.3466
LR&D	-0.039297*	(-3.173529)	0.0020
Lw	-0.126456**	(-2.280360)	0.0249

Note: (*), (**) denotes tests significance at the 1%, and 5% levels respectively

Table16: Panel (FMOLS) estimation for EU-12 countries in high-tech manufacturing sector.

Variables	FMOLS	t-statistic	Prob.
Heterogeneous			
Lgi	0.146448	(1.040984)	0.3014
LR&D	0.756212*	(4.796197)	0.0000
Lw	0.046325	(0.309522)	0.7578
Homogenous			
Lgi	0.207248***	(1.701788)	0.09590
LR&D	0.027917	(0.646444)	0.5204
Lw	0.176395	(0.978595)	0.3316

Note: (*), (**), (***) denotes tests significance at the 1%, 5% and 10% levels respectively. Δ denotes the first difference. L denotes the logarithmic form of the variables

4.4 Conclusion and Policy Recommendation

In the present study, we used annual data for 12 European countries between the periods 1999 and 2009. We employed unit root tests and cointegration tests within the panel framework as well as FMOLS long-run estimation procedures. Tests results of unit root analysis shows that the ll, lR&D, lw and lgi series are integrated at 1st order or follow an I(1) process. Panel cointegration tests which control for a duality of homogeneous and heterogeneous dynamics uncover that the no-cointegration null hypotheses is rejected mostly at the 1% and 5% level of significance. Thus, the presence of a long-run relationship between employment and other exogenous variables is validated. In analyzing the effect of the exogenous variables on the endogenous variable across the different industries. As a result, the long-run elasticity in the full sample is (0.030586) and (0.030486) for both homogenous and heterogenous panels respectively, which is inelastic but significant, the coefficient sign being consistent with our expectations. The coefficient signs for lgi and lR&D are positive and for lw is negative, which is consistent with expectations for our 3 subsamples (services and low-tech and high-tech manufacturing) except wage and capital in low-tech manufacturing that they are insignificant. Shifting our focus towards the primary variables of interest (R&D) and employment. Estimation results from services and high-tech manufacturing are slightly close to those obtained for the full sample of total industry, while there is a negative effect of R&D activity on employment, though insignificant. Innovation does not seem to have any significance for labor demand in the low-tech manufacturing sectors (with losses for gross investments). Generally, the primary discovery made by this chapter of the thesis is straightforward, the employment-friendly component of countries' R&D investments has been revealed to be statistically significant, albeit, with different magnitudes across sectors.

In the present study, our attention has been focused on a primary innovation indicator known as R&D expenditures, although it is strongly related to labor-friendly product innovation, this indicator does not correctly capture the substitute mode of technological innovation, i.e. (possible) employment- inducing process innovation. The underlying implication of this is that the study does not adequately capture the technological change embedded in process innovation with their potential negative impact on employment. Another thing that emerges from the study is that the positive and significant effect of R&D expenditure on employment cannot be observed equally across the services sectors as well as the low and high-tech manufacturing sectors.

By focusing mainly on total employment this study has some limitations. Since a major aim of the European policy agenda is the creation of not just more jobs but better ones, an empirical analysis of inter-sectoral R&D spillovers would be an ideal complement of this study.

Chapter 5

EMPLOYMENT AND OUTBOUND TOURISM DEMAND ELASTICITIES IN OECD COUNTRIES

5.1 Introduction

In recent years, there has been a large influx of studies on tourism development, tourism demand and the tourism-led growth hypothesis. This is largely due to the growing importance of tourism as a major source of revenue and foreign exchange earnings. The tourism demand literature is biased towards studies that analyse inbound tourism demand while neglecting outbound tourism. This may be due to the economic benefits accrued from inbound travel receipts as against the capital flight that is incurred from outbound tourism. It is however important for domestic and international tourism stakeholders to understand the determinants of outbound tourism demand to be well acquainted with external competition. Understanding the determinants of outbound tourism demand at the panel level is also important for policy formulation at the macroeconomic level. This is because at the panel level, outbound tourism demand represents potential inbound tourism demand in other countries both in and outside the panel. In a particular departure country, outbound tourism demand for a specific destination country has also been seen to have strong predictive content for demand for a different destination country (Seo et al., 2010). Thus, stakeholders and policymakers will be better able to understand outbound tourism flow dynamics conditional on the rate of employment and use this information to better predict potential inbound tourist flows to their own economies. This is important because the growth enhancing role of tourism has been well-documented in the literature (Balcilar et al., 2014; Eugenio-Martin et al., 2004; Fahimi et al., 2018; Fayissa et al., 2008; Roudi et al., 2018; Sequeira and Nunes, 2011). Also, assessing the rate at which tourists demand outbound tourism conditional on specific factors in their economies would enable tourism stakeholders develop better marketing frameworks to attract tourists from the various geographical locations in which the studies are undertaken. As of 2016, the world tourism market had a total value of about 1.226 trillion USD with about half of that value (624,750 billion USD) coming from Organization for Economic Co-operation and Development (OECD) countries by way of expenditures and more than half (740,668 billion USD) by way of receipts. The total OECD tourism market maintains about a 4.2% share of gross domestic product (GDP), a 6.9% share of employment and a 21.7% share of services exports in the OECD area (OECD, 2018). Taking into consideration the size of the OECD tourism market, it can be seen why analyzing the determinants of outbound tourism demand in OECD countries would be an important addition to the literature. Studies like the present one would lay the groundwork in the assessment of feasible ways to tap into the large OECD tourism market. Factors that can affect the demand for tourism in positive or negative ways as uncovered in the literature are prices, income, exchange rates, distance and climate (Cheng, 2012; Lin et al., 2015; Salman, 2003). The emphasis of the present study would be on the short- and long-run relationship between employment, income and outbound tourism demand in OECD countries, which is yet to be established in the literature. Also, various studies (Chi, 2016; Dritsakis, 2004; Mun˜oz, 2007; Narayan, 2004; Seetaram, 2012; Smeral and Song, 2015; Surugiu et al., 2011; Untong et al., 2015) have uncovered a positive above unity income effect in the tourism GDP per capita relationship, which implies that tourism is a luxury. This necessitates further studies to assess the income tourism relationship to ascertain if this relationship would still hold when employment effects are controlled for. Our study's contribution to the literature comes in four ways: first, panel studies with a large set of countries which analyse the short- and long-run outbound tourism demand elasticities for the 32 OECD countries employed in the present study are, as at the time of writing, not available in the literature. Second, establishing the existence of a long-run employment and tourism demand relationship would give a more lucid understanding of the relationship amongst income earning employment and expenditure based leisure travel in OECD countries. Third, by employing the Westerlund (2007) and Persyn and Westerlund (2008) bootstrap panel cointegration tests, the Pesaran and Smith (1995) mean group (MG) estimator, the Pesaran (2006) Common Correlated Effects MG (CCEMG) estimator and the Augmented MG (AMG) estimator (Eberhardt and Teal ,2010; Bond and Eberhardt, 2009) , our cointegration test and long-run estimates would be robust to the potential effects of cross-country heterogeneity and cross-sectional dependence. In doing so a clearer and more in-depth picture of the relationships between the variables is elucidated especially as they move through time. Finally, since the long-run coefficients do not show the direction of causality we employ panel Granger causality tests within a panel vector error correction framework to ascertain the direction of causality.

5.2 Data and Model

5.2.1 Data

This study employs annual data for the empirical analysis. All of the data employed for the study come from the same source which is the World Development Indicators (World Bank). The datasets are chosen based on availability for the 32 countries in the panel, within the period 1995 to 2016. International tourism expenditures at current

US$ is used as the only proxy for tourism demand because it has a more complete data set compared to tourist departures. It is also deflated with the US consumer price index to reflect real values. GDP is measured at constant 2010 USD (Halcioglu, 2010; Dogru and Sirakaya-Turk, 2016; Fereidouni et al., 2017). We modify the model of Chi (2016) by including real GDP per adult (ages 15+) as a proxy of real income instead of average wages. This is done by normalizing real GDP on the adult population (population ages 15+). The advantage of using GDP instead of wages is that GDP accounts for income earned by all factors of production including labour. Controlling for only wage earnings in a model with employment may bias the employment estimates. This is because corporate profits in the form of dividends to shareholders (some of whom may not be employed) and investment earnings can also be used to finance outbound tourism. Also, non-wage earnings may constitute a sizable portion of income accrued to senior citizens who may be too old to work but not too old to invest and travel. Furthermore, a decline in the level of the labour income share has been observed in advanced and emerging economies (Dao et al., 2014; Berger and Wolff, 2017) this has been alluded to have come about from technology shocks and the reduction in the cost of investment goods. This scenario may have led to an increased substitution of labour with capital and may also weaken the relationship between employment and outbound tourism demand. A positive sign on the income variable would indicate support for the income effect while a negative one supports the substitution effect. Income coefficient above unity renders support to the hypothesis that tourism is a luxury commodity (Seetaram, 2012; Halcioglu, 2010). A less than unity coefficient supports otherwise (Fereidouni et al., 2017; Dogru and Sirakaya-Turk, 2016). To capture employment levels across OECD countries, the number of employed individuals normalized on adult population (employment to population ratio for over 15 year olds (modeled ILO estimate) is used as a proxy. The analysis is carried out on per capita basis. All variables except relative prices are normalized on adult population (ages 15+). This is done because expenditure on tourist trips is offset by not only the members of the working age population (ages 15-65) but also the senior population (ages 65+). Following Halcioglu (2010), Seetaram (2012) we also employ the trade weighted and effective exchange rate adjusted relative prices to proxy for relative tourism prices. The use of the variable is due to the non-availability of more precise data with sufficient time series that could uniformly capture the relative prices between the departure and destination countries.

The countries included in the study are Australia, Austria, Belgium, Canada, Chile, Czech Republic, Denmark, Finland, France, Germany, Greece, Hungary, Iceland, Ireland, Israel, Italy, Japan, Korea, Latvia, Luxemburg, Mexico, Netherland, New Zealand, Norway, Poland, Portugal, Slovak Republic, Spain, Sweden, Switzerland, United States and United Kingdom. Descriptive statistics for all variables are given in Table 17. It can be observed from Table 17 that employment has the greatest variability in the dataset. The real effective exchange rate has the least spread in the dataset.

Table 17: Summary statistics

Variable	Obs	Mean	Std.Dev.	Min	Max
lritexp	704	55.81157	6.701646	38.648	76.84600
lrgdpp	704	10.7578	.650838	8.9852	12.01798
emp	704	-4.57188	.1458751	-5.11124	-3.880774
lrp	704	6.849648	0.9510212	4.520429	9.382554

Notes: lritexp denotes the natural logarithm of the tourism expenditure in real values normalised on adult population, emp denotes employment to adult population ratio, lrgdpp denotes income (proxied by GDP per capita) and lrp denotes a basket of trade weighted effective exchange rates.

5.2.2 Empirical Model

The long-run relationship between tourism demand and other control variables can be modelled explicitly within a panel framework as,

$$lritexp_{it} = \beta_{0i} + \beta_{1i} lrgdpp_{it} + \beta_{2i} emp_{it} + \beta_{3i} lrp_{it} + \varepsilon_{it} \tag{17}$$

In Eq. (17), lritexp, emp, lrgdpp, and lrp denotes respectively, the natural logarithm of tourism expenditure per adult in real values, employment to adult population ratio, natural log of income (proxied by GDP per adult) and a basket of trade weighted and effective exchange rates adjusted relative prices between departure country and selected foreign trade partners (Halcioglou, 2010; Seetaram, 2012) which is measured as:

$$lrp = \left(\frac{cpi_{fw}}{cpi_d \times ER_{fw}}\right)$$

where cpi_{fw} and cpi_d denotes respectively, the weighted average consumer price index of the foreign trade partners of each departure country and the consumer price index of each departure country in the panel. The CPI has been empirically validated by Morley (1994) to be a veritable proxy for tourism prices due to its high correlation with major tourist expenditure items in 10 important tourist destinations. ER_{fw} denotes the exchange rate between each departure country in the panel and their foreign trade partners. In a particular year, weights are assigned to each foreign trade partner based on the volume of trade it has with the departure country in the same year. Thus, foreign trade partners having higher trade flows with a particular departure country are assigned higher weights. An increase in *lrp* implies that foreign price levels are rising relative to domestic price levels. This variable however has some shortcomings because tourism flows and trade flows can be very different for some countries, nevertheless it serves as the closest proxy to relative tourism prices because of a lack of data to capture tourism weighted relative prices.

Countries are represented by the subscript i (i = 1,…,N) and the time period is indicated via subscript t (t = 1,…,T) . Country explicit effects are denoted by β_{0i} and ε_{it} represents the stochastic error term of country i at time t. Prior to estimating the parameters, the variables are changed to their natural logarithmic form and as such the coefficient estimates can be interpreted as elasticities. The implication of slope heterogeneity is that the model parameters are distinct across countries ($\beta_{1i}, \beta_{2i}, \beta_{3i},$) as against the assumption of slope homogeneity which assumes that model coefficients are identical across countries ($\beta_{1i} = \beta_1, \beta_{2i} = \beta_2, \beta_{3i} = \beta_3,$). If the assumption of slope homogeneity holds, then the usual panel regression methods like pooled OLS (POLS) and fixed effects models can be used to increase the statistical power of the estimation. This comes with the implication of super consistency. Models with slope parameters that are heterogeneous can be estimated employing the mean group (MG) estimators (eg. Pesaran and Smith, 1995; Pesaran et al., 1999) or modifications of mean group estimators.

If all the variables individually follow an I(1) process and the linear combination of all of the variables results in an I(0) process for ε_{it}. Following Pesaran and Smith (1995) Pesaran et al. (1997,1999), Falk (2015), the long-run relationship would then follow an error-correction process of the form;

$$\Delta lritexp_{it} = \alpha_{1i} + \sum_{k=1}^{q} \theta_{1ik}\, \Delta lritexp_{it-k} + \sum_{k=1}^{q} \theta_{2ik}\, \Delta lrp_{it-k} + \sum_{k=1}^{q} \theta_{3ik}\, \Delta lrgdpp_{it-k} + \sum_{k=1}^{q} \theta_{4ik}\, \Delta lemp_{it-k} + \lambda_{1i}(lritexp_{it-1} - \beta_{0i} - \beta_{1i} lrp_{it-1} - \beta_{2i} lrgdpp_{it-1} - \beta_{3i} lemp_{it-1}) + u_{it} \qquad (18)$$

where in the short and long-run values are synchronized by an adjustment mechanism known as the error-correction mechanism or *ecm* (λ_{1i}). The *ecm* is the rate at which short-run deviations are corrected to maintain long-run equilibrium. Adjustment to long-run equilibrium implies λ_{1i} is negative and greater than -1 (no overshooting), which also simultaneously implies the presence of cointegration and long-run predictive content flowing jointly from the exogenous variables to the endogenous one. Eq. (18) would enable the separation of short-run causality from long-run causality via a panel error-correction model. The significance of an θ_{jik} parameters would entail that short-run causality flows from the specific variable to the endogenous variable while the significance of λ_{1i} entails that predictive content flows jointly from the exogenous variables to the endogenous variable at long-run levels.

5.3 Econometric Methods and Results

5.3.1 Cross-Sectional Dependence Tests

Before analyzing the time series properties of the variables in the model we employ the Pesaran (2004) CD test for cross-country correlation in longitudinal data. Cross-sectional dependence in panel data can arise as a result of unobserved common factors, regional or global spatial effects such as the global financial crises of 2007-2010 and/or cross-country spillover effects such as the Asian financial crises in the late 90's. The effect of cross-country correlation can lead to significant size distortions in analyses involving panel unit root and panel cointegration tests of the 1st generation (Baltagi and Pesaran, 2007). Results from Table 18, Panel A, in the CD-test column shows that the statistics for all variables are significant at the 1% level. This rejects the hypothesis for no cross-sectional dependence and shows that all the variables are significantly cross-sectionally dependent. Second generation panel estimation procedures should thus be employed to mitigate the probable size altering effects of cross-sectional dependence.

5.3.2 Panel Unit Root Tests

Before proceeding with the panel analysis, it is necessary to ascertain the time series properties of the variables in the model. Standard conventional panel unit root tests such as the Levin et al. (2002) and the Im et al. (2003) assume cross-sectional independence in the panel, a consequence of which can lead to misleading inferences. In order to circumvent this, we employ the Pesaran (2007) unit root tests for heterogeneous panels with cross-sectional dependence. The test is based on a simple average of the individual augmented Dickey-Fuller (ADF) t-statistics of each cross-section of the panel. The null hypothesis is consistent with non-stationarity of all series. In order to mitigate panel cross-section dependence, cross-sectional averages of lagged levels and first-differences of the individual series are augmented to the standard ADF regressions. Results from Table 18 shows that an I(1) process is followed by all variables.

Following the seminal work of Perron (1989), it has been well established that unit root tests might have low power in the presence of structural breaks. To check the robustness of our results we use the panel stationarity tests of Carrion-i-Silvestre et al. (2005), which controls for incidences of multiple structural breaks. Carrion-i-Silvestre et al. (2005) generalises the KPSS stationarity test of Kwiatkowski et al. (1992) to panel data that allows multiple breaks in both the constant and trend of the model. We estimate the optimal number of constant plus trend breaks using the LWZ information criteria of Bai and Perron (1998). Unit root tests with multiple breaks under the assumptions of both homogenous and heterogenous variance are given in Table 18 Panel B. In both columns the *p*-values are significant at the 1% level for all series. Thus, the null hypothesis of stationarity with multiple breaks is rejected at 1% level for all series. Based on the obtained results becomes appropriate to proceed to the cointegration tests.

Table 18 : Tests for cross-country dependence and unit root tests (Intercept and trend)

Panel A			
Variables	CD-test	CADF Zt-bar stats. (levels)	CADF Zt-bar stats. (1^{st} difference)
lritexp	34.28*	1.687	-11.252*
lrgdpp	90.91*	-0.694	-12.770*
emp	16.00*	3.920	-6.736*
lrp	14.30*	-0.921	-11.874*

Panel B		
	Panel stationarity test with structural breaks and homogeneous variance	Panel stationarity test with structural breaks and heterogeneous variance
lritexp	36.60* (*p*-value: < 0.001; num. breaks: 3)	149.71* (*p*-value: < 0.001; num. breaks: 3)
lrgdpp	42.67* (*p*-value: < 0.001; num. breaks: 4)	251.91* (*p*-value: < 0.001; num. breaks: 4)
emp	46.30* (*p*-value: < 0.001; num. breaks: 4)	625.10* (*p*-value: < 0.001; num. breaks: 5)
lrp	36.20* (*p*-value: < 0.001; num. breaks: 4)	247.10* (*p*-value: < 0.001; num. breaks: 4)

Notes: Based on Authors' calculations. The number of breaks is estimated using the LWZ information criteria of Bai and Perron (1998) allowing for a maximum of $m^{max} = 5$ structural breaks. Structural breaks in both constant and trend are allowed. The long-run variance is estimated using the quadratic spectral kernel with automatic spectral window bandwidth selection as in Andrews (1991), Andrews and Monahan (1992) and Sul et al. (2005). The *p*-values of the panel stationarity tests with multiple structural breaks are obtained with 2,000 bootstrap replications. Significance at the 1% level denoted by *.

5.3.3 Bootstrapped Panel Cointegration Tests

The Westerlund (2007) panel cointegration test proposes four new tests under the null of no cointegration. The tests are based on structural rather than residual dynamics, as a result the common factor restrictions[1] imposed on tests that are grounded on the dynamics of the residual are relaxed as the failure of this restriction can lead to a significant loss of power for these tests (Kremers et al., 1992). The removal of the common factor restriction implies that adjustment processes of both long and short-run are allowed to differ.

The group mean statistics are designed to test the alternative hypothesis that at least one of the cross-sectional units in the panel are cointegrated:

$$G_\tau = \frac{1}{N}\sum_{i=1}^{N}\frac{\hat{\alpha}_i}{SE(\hat{\alpha}_i)}, \qquad G_\alpha = \frac{1}{N}\sum_{i=1}^{N}\frac{T\hat{\alpha}_i}{\hat{\alpha}_i(1)}$$

The above expression implies that the 1st group mean statistics G_τ is obtained by the simple average of N individual t-ratios of least squares estimated country specific error correction parameters denoted as $\hat{\alpha}_i$. This is calculated by normalizing $\hat{\alpha}_i$ on its conventional standard error $SE(\hat{\alpha}_i)$. The second group mean statistics G_α is also obtained by a simple average of N individual t-ratios where in this particular case $\hat{\alpha}_i(1)$ is obtained by estimating $\alpha_1(1) = 1 - \sum_{j=1}^{p_i}\alpha_{ij}$ where $\sum_{j=1}^{p_i}\alpha_{ij}$ is obtained from the least squares equation:

$$\Delta y_{it} = \hat{\delta}_t' d_t + \hat{\alpha}_i y_{it-1} + \hat{\lambda}_i' x_{it-1} + \sum_{j=1}^{p_i}\hat{\alpha}_{ij}\,\Delta y_{it-j} + \sum_{j=0}^{p_i}\hat{\gamma}_{ij}\,\Delta x_{it-j} + \hat{e}_{it} \qquad (19)$$

From Eq. (19) $d_t = (1, t)'$ denotes the deterministic components while $\delta_i = (\delta_{1i}, \delta_{2i})'$ is the concomitant vector of parameters. It can be seen that $\hat{\alpha}_{ij}$ and $\hat{\gamma}_{ij}$ are individual lag parameters of Δy_{it-j} and Δx_{it-j}, respectively. $\hat{\alpha}_i$ denotes the error correction parameter. Details of the test procedure can be obtained from Westerlund (2007).

The panel statistics tests the alternative hypothesis of global panel cointegration:

$$P_\tau = \frac{\hat{\alpha}}{SE(\hat{\alpha})}, \qquad P_\alpha = T\hat{\alpha}$$

where in $\hat{\alpha}$ denotes the error correction parameter which is common across countries and its associated standard error $SE(\hat{\alpha})$ and T denotes the number of observations. To mitigate the effects of cross-country correlations, the bootstrapped error-correction statistic is obtained via the bootstrap approach. This makes the results robust to very general forms of cross-sectional dependence. A detailed outline of the bootstrap procedure is given in Westerlund (2007).

From Table 19, the cointegration tests results show that cross-sectional dependence does not have a significant distorting effect on the test procedure because the results from the ordinary *p*-values and the Robust *p*-values are not very different. While the ordinary *p*-values shows the significance of the conventional test results, the robust *p*-values presents the significance of tests obtained by the bootstrapping procedure. The robust and ordinary *p*-values infer the rejection of the null of no cointegration for both the group mean (G_τ) and panel (P_τ, P_α) statistics at the 5% level of significance. Only the G_α statistics is not significant in both columns. This implies strong evidence for the existence of a common error-correction parameter for the whole panel as well as *N* individual group specific error-correction parameters within the panel. This gives a valid support for cointegration.

Table 19: Westerlund ECM panel cointegration test by bootstrapping

Statistics	Values	*p*-values	Robust *p*-values
G_τ	-5.409**	0.026	0.013
G_α	-2.062	0.985	0.234
P_τ	-17.150***	0.000	0.001
P_α	-7.166**	0.005	0.022

Notes: Significance at 5% and 10% denoted by ** and *** and are indicative of the robust *p*-values. Bootstrap procedure employed 2,000 replications.

5.3.4 Panel Estimation Results

Based on the results obtained from the panel cointegration tests, panel specific and cross-section specific estimation is supported for the model. This implies the estimation of both pooled and grouped models is feasible. To ensure consistency and heterogeneity, we employ the Mean Group (MG), the Common Correlated Effects

Mean Group (CCEMG) and the Augmented Mean Group (AMG) estimators of Pesaran and Smith (1995), Pesaran (2006), Eberhardt and Teal (2010) and Bond and Eberhardt (2009), respectively, to obtain the long-run coefficient estimates. All three estimators follow the same basic procedure. They all estimate group specific regressions and obtain the average of the coefficients across groups. However, the CCEMG and the AMG estimators both use different procedures to control for common correlated effects while the MG does not. In the CCEMG procedure the group-specific regressions are augmented with cross-section specific averages of the endogenous and exogenous variables in order to partial out the effects of cross-sectional dependence. The coefficients are averaged across groups and treated as nuisance parameters. In the AMG procedure however, the set of unobservable common factors are modelled within the framework of a production function estimation. The common factors are treated as a total factor productivity (TFP) type process known as a common dynamic process which is potentially meaningful depending on the estimation. We also employ the fixed effects OLS to increase the power of the estimation and to determine if the parameter estimates at the long-run level are robust to a variety of measurement specifications. Here also, we control for cross-sectional dependence by employing standard errors that are not sensitive to different arrays of both cross-country correlation and autocorrelation up to specific lags.

From Table 19a, the results show the importance of accounting for common factors within the panel as the price effects of the CCEMG and the AMG estimates which is the coefficient on *lrp* are quite similar in both short- and long-run equations implying that demand for outbound tourism in OECD countries is price inelastic. The MG estimation however indicates an elastic price effect in both the short and the long run.

The fixed effects OLS (FE-OLS) estimates in Table 19b also supports a significant inelastic price effect in the long run with evidence for a unitary elastic price effect in the short run, the significance of which is sustained after employing the Driscoll and Kraay (1998) standard errors. The coefficient of employment (*emp*) for the MG, the CCEMG the AMG and the FE-OLS specifications are all insignificant at long run levels but significant in the short run implying that the outbound tourism effect of employment is limited to the short run. The long-run employment coefficient of the FE-OLS estimates loses significance when cross-sectional dependence is controlled for through the Driscoll and Kraay (1998) standard errors. Income effects from Table 4a which is the coefficient on *lrgdpp* show that the AMG and the CCEMG estimators both support a unitary elastic income effect in the long run. However, while the CCEMG consistently supports unitary elastic income effects in the short run, the AMG estimate supports an inelastic income effect in the short run. Income effect for the MG estimate is elastic at short run levels but inelastic at long run levels. For the FE-OLS estimation an inelastic income effect is supported in both short and long run equations. When cross-sectional dependence is accounted for (AMG, CCEMG and FE-OLS with Driscoll and Kraay standard errors), the coefficient estimates seem to be fairly robust across different estimation techniques in terms of signs and statistical significance. Also, the disparities in coefficient magnitudes are fairly minimal showing that the estimated coefficients are not too sensitive to different estimation techniques. However, in order to obtain a single voice in terms of demand elasticities, the Hausman test is employed to ascertain the preferred estimator. Under the null hypothesis (H_0) of the Hausman test, there is no systematic difference between the designated efficient estimator and the designated consistent estimator. Non rejection of H_0 implies that the designated consistent estimator is consistent but the designated efficient estimator is both efficient and consistent and thus is the preferred estimator. Rejection of H_0 however implies that the designated efficient estimator is inconsistent which makes the consistent estimator the preferred estimator. From Table 19c it can be inferred that while the AMG is preferred to the CCEMG estimator, the FE-OLS estimator is preferred to both the CCEMG and AMG estimators as can be seen from the non-rejection of the null hypothesis in all Hausman specification tests.

With this in mind, it can be inferred from the FE-OLS results that tourism is a normal activity in the OECD area which is consistent with Dogru and Sirakaya-Turk (2018) in the Turkish case. Also, the price elasticity of outbound tourism in OECD countries is unity at short run levels and less than unity at long run levels. This shows that relative price appreciation in destination countries will not lead to the accrual or loss of revenues for destination countries in the short run and would generate a net-revenue effect for destination countries in the long run.

Table 19: Mean group estimations (Heterogenous slopes and intercepts)

Variables (Level)	MG	CCEMG	AMG
lrgdpp	0.7916*	1.1390*	0.9822*
emp	0.0081	0.0132	0.0155
lrp	-1.5635*	-0.8835*	-0.8304*
Variables (1st difference)			
Δlrgdpp	1.2818*	1.0073*	0.7457*
Δemp	0.0173*	0.0180*	0.0212*
Δlrp	-1.2965*	-0.7957*	-0.6830*

Notes: Based on Authors' calculations. Significance at the 1%, 5% and 10% levels are denoted by *, ** and *** respectively.

Table 20: Panel fixed effects estimation (Homogenous slopes and heterogenous intercepts)

	Coefficients	
Variables (Level)	With asymptotic S.E	With Driscoll and Kraay S.E
lrgdpp	0.7002*	0.7002*

emp	0.0140*	0.0140
lrp	-0.7821*	-0.7821*
Variables (1st difference)		
Δlrgdpp	0.8867*	0.8867*
Δemp	0.0221*	0.0221**
Δlrp	-1.0600*	-1.0600*

Notes: Significance at the 1% and 5% level is denoted by * and** respectively.

Table 21: Hausman Test

Consistent estimator	CCEMG	CCEMG	AMG
Efficient estimator	AMG	FE-OLS	FE-OLS
χ^2 (Prob > χ^2)	0.61 (0.8952)	1.27 (0.7372)	1.29 (0.7315)

Notes: Under the null hypothesis the efficient estimator is the preferred estimator.

5.3.5 Panel Granger Causality Tests

The existence of cointegration among the variables implies the existence of causality in at least one direction. In order to uncover the latent causal dynamics embedded in Eq. (17) we relax the restrictions imposed on the single equation model in Eq. (18) by assuming that all the variables are endogenous within a panel vector error-correction framework of the following form (Peseran et al., 2009):

$$\Delta lritexp_{it} = \alpha_{1i} + \sum_{k=1}^{q} \theta_{11ik}\, \Delta lritexp_{it-k} + \sum_{k=1}^{q} \theta_{12ik}\, \Delta lrp_{it-k} + \sum_{k=1}^{q} \theta_{13ik}\, \Delta lrgdpp_{it-k} + \sum_{k=1}^{q} \theta_{14ik}\, \Delta emp_{it-k} + \lambda_{1i}\varepsilon_{it-1} + u_{1it} \quad (20)$$

$$\Delta lrp_{it} = \alpha_{2i} + \sum_{k=1}^{q} \theta_{21ik}\, \Delta lritexp_{it-k} + \sum_{k=1}^{q} \theta_{22ik}\, \Delta lrp_{it-k} + \sum_{k=1}^{q} \theta_{23ik}\, \Delta lrgdpp_{it-k} + \sum_{k=1}^{q} \theta_{24ik}\, \Delta emp_{it-k} + \lambda_{2i}\varepsilon_{it-1} + u_{2it} \quad (21)$$

$$\Delta lemp_{it} = \alpha_{3i} + \sum_{k=1}^{q} \theta_{31ik}\, \Delta lritexp_{it-k} + \sum_{k=1}^{q} \theta_{32ik}\, \Delta lrp_{it-k} + \sum_{k=1}^{q} \theta_{33ik}\, \Delta lrgdpp_{it-k} + \sum_{k=1}^{q} \theta_{34ik}\, \Delta emp_{it-k} + \lambda_{3i}\varepsilon_{it-1} + u_{3it} \quad (22)$$

$$\Delta lrgdpp_{it} = \alpha_{4i} + \sum_{k=1}^{q} \theta_{41ik}\, \Delta lritexp_{it-k} + \sum_{k=1}^{q} \theta_{42ik}\, \Delta lrp_{it-k} + \sum_{k=1}^{q} \theta_{43ik}\, \Delta lrgdpp_{it-k} + \sum_{k=1}^{q} \theta_{44ik}\, \Delta emp_{it-k} + \lambda_{4i}\varepsilon_{it-1} + u_{4it} \quad (23)$$

From Eq. (20) to (23), short-run causality is validated by the joint rejection of the null hypothesis; H_0: $\sum_{k=1}^{q} \theta_{jik} = 0$ for $k = 1,\ldots,q$ lags of the i-th variable in the j-th equation. Long-run causality is conducted by testing H_0: $\lambda_{ji} = 0$ for all i. The coefficient λ of ε measures how fast the deviations from the long-run equilibrium are readjusted back towards the equilibrium path following changes in the level of each variable. The causality tests are calculated following the dynamic heterogeneous panel approach of Peseran et al. (2009). From Table 20 short-run causality is validated through the significance ($p < 0.5$) of the χ^2 statistics of the causing variables located in the second to fourth columns of the table. It is thus validated that in the short run, bidirectional causality exists between tourism expenditure and exchange rate adjusted relative prices. A short-run bidirectional relationship may stem from the fact that tourism expenditure can contribute to capital flight in departure countries which has a tendency to increase interest rates and instigate currency depreciation in those countries. Once the currency depreciates it becomes more expensive to undertake outbound tourism from these countries and less expensive for foreign countries to undertake inbound travels to same countries. This brings about a net reduction in outbound tourism expenditure from departure countries and net increase in revenue from inbound travels. When the revenue from inbound travels increases sufficiently enough, currencies from the departure (domestic) countries may begin to appreciate relative to the foreign countries due to the movement of capital from foreign countries to departure (domestic) countries. These effects may not be significant when empirically analysed with country specific data because the size of tourism relative to the economy of each OECD country may be quite small. They can however appear sizable when panel studies are considered due to the accumulation of different country specific effects. Bidirectional causality exists between income and employment while unidirectional predictive content flows from income to tourism expenditure which is in consonance with Halcioglu (2010).

In the long-run scenario, long-run causality is a mutual effect amongst all the variables in the system. In order to compare speed of adjustment, we us the half-life of the shock under homogeneity assumption, which is calculated as $\ln(0.5)/\ln(1-|\lambda|)$ where λ is the coefficient on the error-correction term under homogeneity assumption $\lambda = \lambda_i$. The error-correction terms indicate a slower adjustment rate to long-run equilibrium for relative prices and tourism expenditure with income and employment having faster adjustment speeds, with half-life of the shocks in about 1.3 years for tourism expenditures while that of relative prices attains half-life in about 1.9 years. Half-life of the shock to employment is about 2.9 years while half-life is attained in about 6.8 years for income. The result implies that income and employment are the most endogenous in the system. Only income and price have short-run predictive content for tourism expenditure. Employment may however affect tourism expenditure via its impact on income.

Table 20: Granger causality analysis. (Panel vector error-correction)

Endogenous variables		← Causal flow (Causing variables)				
		Short- run				Long run
		Δlritexp	Δemp	Δlrgdpp	Δlrp	ECT_{t-1}
Eq. (17)	Δlritexp	---	1.85	7.66^{**}	7.43^{**}	-0.41^{*}
Eq. (18)	Δemp	4.30	---	23.56^{*}	3.31	-0.21^{*}
Eq. (19)	Δlrgdpp	4.08	11.64^{*}	---	3.67	-0.10^{*}
Eq. (20)	Δlrp	16.36^{*}	0.51	0.44	---	-0.31^{*}

Notes: ECT denotes the error-correction parameter. Significance at the 1% and 5% levels are denoted by * and **. Figures in cells labelled 'short-run' denote the χ^2 statistics for the Wald tests of the null H_0: $\sum_{k=1}^{q} \theta_{jik} = 0$. Numbers in the cells labelled long-run indicate the estimated adjustment parameter λ_j under homogeneity assumption $\lambda = \lambda_i$. A lag order of 1 is employed for the estimation based on the AIC and SBIC criterion.

5.4 Conclusion

The present study employs second generation panel cointegration, panel estimation and panel unit root testing techniques to ascertain the relationship between tourism demand employment and real income at long and short-run levels. The Pesaran (2007) unit root tests with cross-sectional dependence uncovered that all the variables followed an I(1) process. Bootstrap panel cointegration analysis supports the presence of error-correction for the whole panel as well as for $n \geq 1$ cross-sectional unit(s) within the panel. Panel estimation with the mean group, the augmented mean group and the common correlated effects mean group estimation procedures show that the demand for tourism is unaffected by employment in the long run but is however positively affected by employment in the short run. Going by the preferred estimator, income elasticity of tourism demand is shown to be inelastic in the short and the long run. The implication of this being that outbound tourism is a normal consumption in OECD countries. Causal analysis via the vector error-correction model within a panel framework shows that short-run unidirectional causality flows from income to tourism expenditure. Also, bidirectional causality exists between relative prices and tourism expenditure.

The fact that the effects of increase in employment and income are both positive at short run levels while only the effect of income persists to the long run may imply rising inequality in income distribution among the employed population. The reason for this may be due to the increment of the capital labour ratio of the production process which has resulted in a reduction in the wage share of GDP and increased corporate savings from the 1980's (Karabarbounis and Neiman 2013). This brings about a scenario whereby the quality of employment as regards to remuneration rather than the quantity of the employed becomes an important determinant of long-run outbound tourism demand. The present study has shown that in the development of marketing and advertisement frameworks for tourism destinations, the overall state of the economy of potential departure countries should be taken into consideration as GDP per capita has been shown to be the better predictor of outbound tourism demand rather than employment. One other notable aspect is the change in the demographics of EU countries and other advanced nations which make up the OECD. An ageing population and increased life expectancy has resulted in the increased share of senior tourists in the tourism market. This may have contributed to the dampening of the long-run significance of employment for outbound tourism. Senior citizens who are above the legally mandated working age (65 in most countries) have a variety of non-wage income sources at their disposal which may be used to offset travel expenditure. Some of these income sources are; occupational as well as state pension schemes, savings and maturing endowment policies as well as an ability to easily re-allocate resources to leisure activities due to reduced financial commitments (Avcikurt, 2009). Tourist stakeholders should not neglect older vacationers when developing their marketing frameworks because this segment of the population constitutes a sizeable market share. The panel estimates indicate that an increase in employment level is significantly related to increased demand for outbound tourism in only the short run. Likewise, income is a significant positive determinant of outbound tourism in both the long run and the short run. Granger causality tests at both long and short-run levels also show that outbound tourism effects of employment are not consistent across time. Also, due to the short-run effect of rising employment rates to tourism spending there is a need to develop tourist programs oriented towards the vacation needs of the employed individuals within the countries under study. Yearly employment rates can serve as indicators of potential tourism receipts from these countries. Also, marketing and promotional programs of potential destination countries can be constantly reviewed to meet up with periodic changes in median wage rates in relationship with destination costs across OECD countries.

Chapter 6

CONCLUSION

The present thesis analyses the determinants and effects of employment in selected regions with particular emphasis on innovation and tourism. To this end, the thesis was divided into 3 self-contained chapters. The first chapter analyses the innovation and employment nexus in 8 Asian economies. Panel cointegration and estimation analysis were employed to determine if the variables commove in the long-run. A long-run relationship was uncovered and thus panel estimation procedures were utilized to determine the long-run cointegrating coefficients. The obtained results show that innovation demonstrates a non-linear relationship with employment. The job destroying effect of innovation becomes evident at lower levels of innovation. However, as the scale of innovation increases, its job creating component begins to counterbalance its job displacing component due to the existence of more innovations and the demand and supply pressures around the industries created by such innovations. Demand pressures for a particular innovation would necessitate employment in the production firms of such innovations. Furthermore, such innovations would potentially create new industries around their use and other logistics which could spur employment.

In the fourth chapter of this thesis the relationship between innovation and intersectoral employment is analyzed. Panel unit roots and cointegration tests uncover a long run relationship amongst the variables. Long-run parameter estimates show that while the innovation job inducing impact is evident in the service and high-tech manufacturing sector, a job destroying effect is however uncovered for the low-tech manufacturing sector. The job destroying effect of innovation in the low-tech manufacturing sector may be due to a possible higher labor intensity relative to the high-tech manufacturing sector. As such any innovation may lead to an increased mechanization which could possibly reduce the need for labor. As such product innovation in the low-tech manufacturing sector may have a higher component of process innovation. In the high-tech manufacturing sector which may be less labor intensive, increased product innovation may not lead to the reduction of labor intensity due to the already highly mechanized nature of the high-tech manufacturing sector. It may however lead to the creation of parallel machines with similar but more sophisticated and simpler procedures. These machines may increase the speed of production, reduce the price of products and expand the workforce due to rising product demand.

In the fifth Chapter of this thesis, the outbound tourism demand effects of employment were analyzed for 32 OECD countries between the periods 1995 to 2016. Panel cointegration which are robust to cross-sectional dependence were utilized to ascertain the long-run relationship amongst the model. Panel estimation techniques which are also robust to cross-sectional dependence are employed to ascertain their short and long-run coefficient estimates. The results demonstrate that employment does not affect outbound tourism in the long-run but has a short-run effect. Panel granger causality tests show evidence of long-run causality running towards all the variables. A short-run uni-directional causality flowing from income to out-bound tourism show that income is the better predictor of outbound tourism demand. However, the importance of employment is predicated on a bi-directional causal relationship between income and employment, thus it is safe to assume that employment affects tourism through its effect on income.

This thesis has several implications from the perspective of supply side as well as demand side policies. On the supply side, stake holders can initiate policies that would benefit the innovation of products with the highest potential of generating backward and forward industries which could generate the highest employment. Grants and subsidies should be awarded for such innovations once it can effectively be proven by demonstrations, feasibility studies and simulations that they have the potential of inducing high employment effects. By doing so, product innovators would critically consider the implications for employment generation when developing innovations which have potential implications for the job market.

For demand side policies, OECD countries should consider the use of annual employment levels to gauge potential increments of outbound tourism expenditure in order to implement short-run taxes. So that the gains from employment will not be totally lost by outbound expenditure. Also, thresholds can be set above which excessive outbound tourism could be taxed.

As much as the thesis has implications for departure countries, it also portends some implications for destination countries. Destination countries which have a large portion of their inbound tourists arriving from OECD countries should consider revising their marketing strategies to reflect OECD employment rates. This could be done by initiating pricing policies which tie tourism prices such as hotel costs to the state of the OECD economies. Putting the results of the two thesis chapters together it can be inferred that wealth effects from high scale innovations could spur outbound tourism demand. This is because since high scale innovations positively induces employment it can then be safe to deduce that its employment creating effect would spillover to the demand for outbound tourism. However, this assertion needs empirical evidence and thus should form the basis for future research because the two separate research where conducted for two different regions composed of different countries.

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