ESSAYS ON THE MACROECONOMICS OF GENDER EQUALITY IN THE LABOR MARKET

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ABSTRACT

ESSAYS ON THE MACROECONOMICS OF GENDER EQUALITY IN THE LABOR MARKET

Kılınç, Dilara

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This thesis provides four essays regarding the macroeconomic implications of the two-way relationship between gender equality in the labor market and income dynamics. Particularly, in the first essay, the role of gender equality on income convergence is investigated for OECD countries within the framework of neoclassical growth theory. Given that any social development indicator that has a high level of interaction with income may also mimic convergence behavior, the research focus of the subsequent essay is on the identification of whether this group of countries will also converge in gender equality measures. The convergence analyses are performed by the GMM estimators to overcome potential modeling problems in dynamic panel data. The findings show strong evidence of the positive contribution of gender equality on income convergence and of absolute and conditional gender equality convergence across OECD countries. In the next essay, the impact of economic development on gender equality is analyzed in the context of 'Gender Kuznets Curve' hypothesis. The estimations by CCE and AMG methods ascertain that though the panel of OECD countries confirms the hypothesis, the country-specific impacts vary and even contradict. Finally, in the fourth essay, time series analyses are provided for G7 countries to determine the stationarity and the long-run income elasticity of gender equality, given the structural breaks in the process. The results by NP unit root test and ARDL approach reveal the shock persistence in gender equality process in all and the positive impact of income on gender equality in five out of G7 countries.

Keywords: Gender Equality; Labor Force; Employment; Aggregate Output; Economic Growth; Convergence; Models with Panel Data; Time Series Analyses

ÖZET

İŞGÜCÜ PİYASASINDA CİNSİYET EŞİTLİĞİNİN MAKROEKONOMİK UYGULAMALARI ÜZERİNE MAKALELER

Kılınç, Dilara

Ekonomi Doktora Programı

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Bu tez işgücü piyasasında cinsiyet eşitliği ile gelir dinamikleri arasındaki çift yönlü ilişkinin makroekonomik çıkarımları hususunda dört makale sunmaktadır. Ayrıntılı olarak, ilk makalede, cinsiyet eşitliğinin gelir yakınsaması üzerindeki etkisi OECD ülkeleri için neoklasik büyüme teorisi çerçevesinde araştırılmıştır. Gelir ile yüksek derecede etkileşim halinde olan sosyal kalkınma göstergelerinin de yakınsama davranışını örnek alabileceğini düşünürsek, bir sonraki makalenin araştırma konusu bu ülke grubunun cinsiyet eşitliği ölçütlerinde de yakınsama davranışı gösterip göstermeyeceğini belirlemek üzerine olmuştur. Dinamik panel veride oluşabilecek olası modelleme sorunlarını gidermek amacıyla yakınsama analizleri GMM tahmincileri tarafından gerçekleştirilmiştir. Bulgular cinsiyet eşitliğinin gelir yakınsaması üzerindeki olumlu etkisini ve cinsiyet eşitliği ölçütlerinin de koşulsuz ve koşullu yakınsama davranışı gösterdiğini OECD ülkeleri için güçlü bir şekilde kanıtlamıştır. Bir sonraki makalede, ekonomik kalkınmanın cinsiyet eşitliğine olan etkisi 'Cinsiyet Kuznets Eğrisi' hipotezi çerçevesinde analiz edilmiştir. CCE ve AMG yöntemleri kullanılarak OECD ülkeleri için yapılan tahminler panel genelinin hipotezi doğruladığını, fakat ülkeye özgü etkilerin çeşitlilik gösterdiğini, hatta hipotez ile çeliştiğini göstermektedir. Son olarak, dördüncü makalede, süreç içerisinde gerçekleşen yapısal kırılmalar da göz önünde bulundurularak cinsiyet eşitliğinin durağanlık özelliği ve uzun-dönem gelir elastikiyeti G7 ülkeleri için zaman serisi analizleri ile incelenmiştir. NP birim kök testi sonuçları şokların cinsiyet eşitliği sürecinde kalıcı etkiler bıraktığını tüm ülkeler için gösterirken, ARDL yaklaşımı tahminleri gelirin cinsiyet eşitliği üzerindeki olumlu etkisini G7 ülkelerinin beşinde sunmaktadır.

Anahtar Kelimeler: Cinsiyet Eşitliği; İşgücü; İstihdam; Toplam Çıktı; Ekonomik Büyüme; Yakınsama; Panel Veri Modelleri; Zaman Serisi Analizleri

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LIST OF ABBREVIATIONS

ADF : Augmented Dickey-Fuller
AMG : Augmented Mean Group
ARDL : Autoregressive Distributed Lag
CADF : Cross-Sectionally Augmented ADF
CCEMG : Common Correlated Effects Mean Group
CIPS : Cross-Sectionally Augmented Im-Pesaran-Shin
CD : Cross-Sectional Dependence
CUSUM : Cumulative Sum
CUSUMSQ : Cumulative Sum of Squares
EU : European Union
FLFPR : Female Labor Force Participation Rate
FSE : Female Share in Employment
GDP : Gross Domestic Product
GKC : Gender Kuznets Curve
GMM : Generalized Method of Moments
G7 : The Group of Seven
ICT : Information and Communications Technology
ILO : International Labour Organization
IMF : International Monetary Fund
LHS : Left Hand Side
MDGs : Millennium Development Goals
NP : Narayan and Popp
OECD : The Organization for Economic Cooperation and Development
RHS : Right Hand Side
SPO : State Planning Organization of Turkey
UN : United Nations

CHAPTER 1

INTRODUCTION

> …investing in women is not only the right thing to do. It is the smart thing to do. I am deeply convinced that, in women, the world has at its disposal the most significant and yet largely untapped potential for development and peace. Gender equality is not only a goal in itself, but a prerequisite for reaching all the other international development goals, including the Millennium Development Goals. (Ban Ki-Moon, UN Secretary General, United Nations Headquarters of International Women's Day on 8 March 2008).

1.1 Motivation

Gender equality implies that women and men have the same status and equal opportunities in all sectors of the society, and that they have equal rights and conditions to contribute to economic, social and political development, and to benefit from the outcomes. Gender equality and women's empowerment have been the third of the eight goals of the global action plan to achieve by 2015 that is concluded by the 2010 Millennium Development Goals (MDGs) Summit. Additionally, the summit issued a call for the adoption of a resolution to assure gender equality in endowments, economic opportunities and agency through gender mainstreaming in development policy making. According to the global action plan, gender equality is a development objective in its own right, besides serving as an instrument for the achievement of other development goals (The MDGs Report of UN, 2010).

The nexus between gender equality and development has been widely discussed under two sets of arguments. The first set of arguments states that gender equality is of importance intrinsically, and is inarguably valued as an end in itself, since development is a process of increasing rights and opportunities equally (Sen, 1999; Kabeer, 2005). Gender equality is in and of itself a core development objective, since the accumulation of physical and human capital endowments to perform labor and to produce an economic value, the utilization of those endowments to take the economic opportunities and to earn incomes, and the application of those endowments to take action in accordance with one's own preference as fundamental rights and opportunities should be equal for everyone, irrespective of whether one is a woman or a man. The second set of arguments asserts that gender equality is a matter of great importance instrumentally, since it serves as a smart development policy. Gender equality in endowments, economic opportunities and agency improves economic efficiency, promotes overall productivity and hence helps ensure income growth and sustainable economic development.

Enhancing women's absolute and relative status contributes to economic development through three channels. First, equalizing educational levels and economic opportunities of women with those of men leads to overall productivity gains in an economy, which become crucial in an ever competitive and globalized world. Gender discrimination in societal institutions and in the labor market that excludes women from educational opportunities, high skilled jobs and well paid work due to their socially constructed and historically developed roles causes the misallocation of skills and talents and high economic losses. The well documented evidence also shows that the inefficiencies in the allocation of inputs due to gender inequalities caused by patriarchal system or by discrimination comes at high economic costs. The recent study of Cuberes and Teignier-Baqué (2011) found that income would be 7%-18% higher across a range of countries if women had the opportunity to work in the same sectors, jobs and positions as men. Similarly, the findings of Löfström (2009) suggest that the average EU income would increase by almost 30% if women worked on the same terms as men. According to Blackden et al. (2007), gender inequality in education reduces human capital accumulation in a society and hinders economic efficiency and productivity due to the restriction of the pool of skills and talents to draw from. The findings of Esteve-Volart (2004) assert that occupational segregation by gender distorts the allocation of talent across occupations and reduces average skill and talent in the workforce, which in turn has negative impacts on economic development. According to her findings, the exclusion of women from high skilled jobs in the labor market causes the reduction in human capital investment for both women and men workers, which results in lower aggregate productivity and lower income. The gender inequality in access to productive inputs distorts the allocation of resources, which reduces aggregate current productivity and investment in new technologies (Sen, 1990; Klasen and Wink, 2003). Using cross-country and panel regressions, the findings of Klasen (1999) assert that gender inequality in education and employment have direct negative impacts on economic growth through lowering the average quality of human capital. The average innate ability in case of favoring educational opportunities of boys is lower than would be the case if boys and girls have equal opportunities and conditions for the access to education. This reduces the average level of human capital in the economy and thus obstructs economic growth (Dollar and Gatti, 1999; Klasen, 2002; Knowles, Lorgelly and Owen, 2002). Gender discrimination induces higher economic losses in a county that is more open to international trade, since it hampers the efficient use of inputs mattering for the country's globalization and economic growth (Do, Levchenko and Raddatz, 2011). Gender inequality hinders the international competitiveness of a country, particularly if it is an exporter of goods and services for which the workforce of woman and men are equally well suited. The evidence also suggests that countries in which women workers are empowered experience larger export shares in relatively female labor intensive sectors (Do et al., 2011). Hence, the countries that get over gender inequality in educational enrolment and economic participation gain a competitive advantage over those that retard the action in a globalized world (World Development Report of The World Bank, 2012). Second, women's human capital investment and economic empowerment enhance societal and economic well-being of the next generation and hence foster future economic growth. Gender inequality in paid employment reduces the bargaining power of women at home that may lead to lower investments in human capital of the next generation (Sen, 1990; Klasen and Wink, 2002). Third, gender equality in economic, social and political activities, decision making and policy making produces more representative and more inclusive institutions and policy choices and thus a sustainable development path in a country

(World Development Report of The World Bank, 2012). The empowerment of women's decision making in institutions promotes aggregate productivity and macroeconomic performance due to their stronger preferences for savings and higher tendencies towards investment in productive projects (Stotsky, 2006).

In the reverse direction, economic development also has a positive causal impact on gender equality. This impact is discussed under two strands of arguments in the literature. The first strand of arguments asserts the linear increasing trend of gender equality in response to economic development, which is achieved through three channels. First, economic development will support societal integration, provide new incentives to invest in human capital and create a wider range of employment opportunities for women, thus will enhance their status in the labor market (Weiss, Ramirez and Tracy, 1976; Clark, 1991; Clark, Ramsbey and Adler, 1991; Charles, 1992; Forsythe, Korzeniewicz and Durrant, 2000; Duflo, 2012). Higher labor market participation of women will increase their bargaining power in the economy, which in turn improves gender equality. Second, economic growth promotes women's status and undermines gender gap in the labor market, since gender discrimination will impose additional costs for institutions in the process of market competition and globalization (Forsythe et al., 2000). Third, economic development brings along the expansion of activities in the service sector that is well suited to female employment and leads to greater participation of women in the wage labor market. Finally, better access to public services in the course of economic development improves human capital accumulation of women and thus reduces gender gap in educational attainment and labor force participation (World Development Report of The World Bank, 2012). The second strand of arguments suggests the nonlinear U-shaped trend of gender equality in response to economic development. Since the pioneering study of Boserup (1970), several studies have argued that the initial stages of economic development are characterized by a decreasing gender equality, which begins to increase once the country develops beyond a certain threshold (Psacharopoulos and Tzannatos, 1989; Goldin, 1995; Tam, 2011). The explanation of this stylized nonlinear behavior of gender equality during development is follows: women are engaged in agricultural activities in large numbers prior to urbanization and industrialization process. In the early stages of industrialization, with the shift of economic activities from agriculture to manufacturing in which men's workforce comes into prominence due to patriarchal social structure and discriminatory institutions, women lose out on the utilization of their skills and talents, and gender equality decreases. Eventually, the later stages of economic development brings along the change in market structure such as the expansion in service activities, which facilitates the accumulation and utilization of human capital for women and encourage them to participate in the wage labor market.

Gender equality in labor market participation has been the key issue to evaluate the extent to which women and men have equal opportunities in the utilization of physical and human capital endowments and in the access to economic activities. Greater participation of women in the labor market provides them an independent income stream, which in turn enhances their bargaining power in the economy, and lends them greater social and economic visibility (Eastin and Prakash, 2013). Over the past quarter century, women have joined the labor market in increasing numbers, partially narrowing the gender gap in participation. In the period 1980-2009, the global labor force participation rate increased from 50.2% to 51.8% for females, while it decreased from 82% to 77.7% for males. As a consequence, gender gap in labor force participation rate decreased from 32 percentage points to 26 percentage points in this period (World Development Report of The World Bank, 2012). The change in the labor market structure and in the nature of institutions, greater economic integration of countries and moderation in patriarchal social norms in conjunction with economic development go a long way towards greater participation of females in the labor market over the past 25 years. Particularly, economic development has brought along growing economic opportunities for women, especially in the service sector. Female workers are more likely to participate in services than male workers, that is, employment in services constitutes 47% of total employment for women, while the share of employment in services in total employment is 40% for men (World Development Report of The World Bank, 2012). Additionally, greater trade openness, competitiveness and economic integration due to globalization have opened the way to the female workforce in export-oriented and ICT-enabled sectors, particularly in developed countries. The competitiveness of countries has also led to institutional changes in the market in such a way that the firms have tended to demand for more part-time or irregular workers with lower wages, which is more suited to females due to their demand for flexible working-time arrangements in general (Standing, 1999). Finally, the change in social norms through rising levels of education, delayed marriage and lower fertility rates have facilitated women's participation in the wage labor market (Goldin, 1990). In view of the fact that women represent at least one half of the potential talent throughout the world, achieving the best utilization of women's skills and qualifications will enhance societal well-being, reduce poverty and boost economic development at national and global levels in the form of increasing investment in human capital (OECD, 2008).

1.2 Organization

The main focus of this thesis is the two-way relationship between gender equality in the labor market and income dynamics. Particularly, it provides four essays of which one concerns the impact of gender equality on income convergence, one examines the convergence behaviors of gender equality indicators across countries, and the other two essays present the different approximations to the impact of income dynamics on gender equality. In gender literature, the female labor force participation rate (FLFPR) is frequently considered as the key instrument for women's status and gender equality, since it indicates economic and social empowerment of the labor market (Eastin and Prakash, 2013). Since the early studies of Mincer (1962) and Cain (1966), several studies have used the female labor force participation as a measure of women's labor market activity, including Hill (1983), Psacharopoulos and Tzannatos (1989), Clark et al. (1991) and Eastin and Prakash (2013). However, this thesis asserts that FLFPR may be inaccurate to report women's overall production efforts, thus may be inadequate to reflect women's active engagement and to represent gender equality in the labor market, since the composition of female labor force also comprises unemployed women who are actively searching for a job in the market. In this regard, this thesis utilizes the female share in employment (FSE) as an alternative indicator of the labor market gender equality.[1] The underlying motivation to use FSE is that this measure only counts in women who are participated in the labor market to produce goods or services, and thus reflects women's status in the labor market and their access to economic opportunities more accurately.

The conditional income convergence hypothesis supposes that countries with similar structural characteristics converge to similar balanced growth paths in terms of income per capita in the long run. Hence, the model can be taken in a broader perspective, for example, in which a social characteristic of a country is included. Starting from this point of view, Chapter 2 investigates the role of the labor market gender equality on conditional income convergence both theoretically and empirically in the framework of Solow-Swan neoclassical growth model. The theoretical part develops a novel approach by augmenting the textbook model with the labor market gender equality. The theoretical model contributes to the literature by determining a 'golden rate' that would imply the absolute labor market gender equality, and that maximizes income per worker. In the empirical part, dynamic panel data approach is utilized, and System Generalized Method of Moments (GMM) estimator is applied to 5-year span unbalanced panel data of 34 OECD countries over the period 1951-2010. The motivation to favor OECD countries for the analysis of the role of the labor market gender equality on income convergence is twofold. First, a substantial share of economic growth in this area comes from increasing participation of women in the labor market in recent decades (OECD, 2008). Second, the narrowing of the gender gap in labor market participation has been a common characteristic in the last few decades due to socio-cultural, economic and institutional improvements across these countries (OECD Employment Outlook, 2002).

At first step, the textbook model is estimated. Next, the impact of the labor market gender equality on income convergence is analyzed within the augmented model. Finally, three control variables, namely trade openness, foreign direct investment

[1] The female share in employment is the ratio of female employment to total employment (aged 15-64). Employment comprises those who are in paid employment and those who are in self-employment (unpaid family workers). The female labor force participation rate is the ratio of female labor force (employed and unemployed, but looking for work) to female working-age population (aged 15-64).

(FDI) inward stock and income inequality are added to the augmented model to increase the explanatory power of the labor market gender equality on income convergence. The System GMM findings of both the textbook model and the augmented model reveal a strong evidence of GDP per worker convergence across OECD countries in the period 1951-2010. The labor market gender equality that is measured by FSE is positive and statistically significant in all augmented model estimations, which confirms the theoretical expectation. Additionally, all control variables are found to be statistically significant and to have expected signs when included one by one in the augmented model, showing that each of them strengthens the relationship between FSE and income convergence. The estimation of the implied speed of convergence in the augmented model is higher than that in the textbook model, which suggests that FSE has a substantial contribution on GDP per worker convergence. Moreover, this contribution is found to be higher when the control variables are included in the augmented model. On the whole, Chapter 2 instantiates that women's empowerment in the labor market have promoted the income convergence across OECD countries in the last five decades.

Given that any social development indicator may also portray a convergence behavior in case of a high level of mutual interaction with income, Chapter 3 grounds on the empirical examination of whether the labor market gender equality indicators also converge across 34 OECD countries over the period 1971-2010. In accordance with the per capita income convergence equation, a dynamic panel equation is originated to estimate absolute (unconditional) and conditional gender equality convergence in the labor market. The estimation procedure of conditional convergence utilizes various control variables that affect the labor market gender equality, such as GDP per capita, tertiary education, fertility rate, trade openness and FDI inward stock. In the empirical analysis, the Difference and the System GMM estimators are employed for the 5-year span unbalanced panel data. The estimations confirm a priori hypothesis that the labor market gender equality indicators mimic income convergence behavior across OECD countries in the period 1971-2010. In particular, all GMM regressions yield strong evidence of FLFPR and FSE convergence in both absolute and conditional senses. In all conditional convergence estimations, the control variables are statistically significant, and their signs are consistent with the literature. Furthermore, the (implicit) speed of convergence conditional on each of control variables is found to be higher than that of absolute convergence, which suggests that each of them contributes to the labor market gender equality convergence. All in all, Chapter 3 provides strong evidence that gender equality convergence in the labor market, which in and of itself is a core social and economic development objective is achieved by OECD countries significantly.

There are three major motivations to employ either the System GMM or both the Difference and the System GMM estimators in Chapters 2 and 3. First, both the Difference GMM estimator proposed by Arellano and Bond (1991) and the System GMM estimator proposed by Arellano and Bover (1995) and Blundell and Bond (1998) cope with modeling issues such as fixed effects and potential endogeneity of regressors, and provide unbiased estimations in dynamic panel data framework (Bond, Hoeffler and Temple, 2001; Hoeffler, 2002). Second, both of the estimators are highly suggested for convergence equations as they fit well the linear equations with one dynamic dependent variable, additional explanatory variables and fixed effects (Bond et al., 2001; Roodman, 2009b). Third, both of them fit for panel data sets with small time dimension and relatively large cross-sectional dimension. However, the System GMM is more efficient than the Difference GMM estimator due to its additional assumption that the first differences of instruments are uncorrelated with the fixed effects, which allows the inclusion of more instruments in a regression (Roodman, 2009a). Moreover, the Difference GMM tends to yield downward biased estimation of the lagged dependent variable if the series are persistent over time (Blundell and Bond, 1998). The consistency of the GMM estimations in Chapters 2 and 3 is detected through three key diagnostics: Arellano-Bond (1991) test for AR(2) in first differences, Hansen (1982) test of over-identifying restrictions and the rule of thumb.[2] All GMM estimations are found to be consistent and robust in terms of the validity of instruments used by the estimators, and of the expected signs and the significance levels of coefficients of the

[2] The Arellano-Bond test results require lack of AR(2) serial correlation in the first-differenced residuals. The Hansen test assesses whether the instruments are uncorrelated with the error term to ensure instrument validity. Finally, the rule of thumb suggests that the number of instruments should be smaller than or equal to the number of groups in a regression.

determinants of the respective convergence models. Additionally, in Chapter 3, the estimations by the System GMM and the Difference GMM are compared with the Difference-in-Hansen test, which suggests that the System GMM is preferred to the Difference GMM providing the additional moment restrictions hold.[3] In all System GMM estimations, the statistics of Difference-in-Hansen test reveal the validity of additional moment restrictions, suggesting that System GMM is the preferred estimator of the convergence analysis in Chapter 3.

In Chapter 4, the impact of economic development on the labor market gender equality is investigated in the context of "Gender Kuznets Curve" (GKC) hypothesis, cf., Eastin and Prakash (2013). The hypothesis suggests that gender equality portrays a U-shaped pattern in the course of economic development. Nevertheless, the evidence for a single country and/or a panel of countries does not necessarily confirm this argument. In this regard, this empirical research is the first in the literature to investigate the potential heterogeneous behavior of gender equality in response to economic development across a panel of countries. Although the ever-narrowing gender gap is a common labor market progress across OECD countries, there are still variations in the timing and the extent to which the narrowing has occurred, which stem from economic factors and public policies of individual countries. Hence, intuition suggests that the relationship between the labor market gender equality and economic development may not be homogeneous across these countries. In this regard, the aim is to determine the country-specific responses of gender equality to economic development, in addition to the common response for the whole panel of 28 selected OECD countries over the period 1990-2012.[4] The empirical analysis utilizes the Common Correlated Effects Mean Group (CCEMG) estimator proposed by Pesaran (2006) and Augmented Mean Group (AMG) estimator suggested by Eberhardt and Bond (2009), which allow cross-sectional dependence that may result from economic integration of countries and/or common financial, political and social shocks. The panel data findings from 28 countries show that the labor market gender equality that is indicated by FSE portrays a U-shaped trend in response to economic

[3] The Difference-in-Hansen test presents the p-values for the null hypothesis of the validity of additional moment conditions.

[4] As the empirical methodology does not allow unbalanced panel data, the Post-Soviet countries that have limited data availability are excluded from the analysis.

development in the period 1990-2012. Nevertheless, the country-specific findings reveal that the response of FSE to economic development varies and even contradicts GKC hypothesis for some individuals. In particular, the Common Correlated Effects (CCE) and the AMG findings show that 13 countries support the hypothesis, whereas three countries portray the reverse relationship between gender equality and economic development, and contradict the hypothesis. Additionally, the estimations for 12 countries imply no statistically significant relationship between FSE and economic development. The empirical findings in Chapter 4 indicate that the labor market gender equality process may display a heterogeneous behavior across individuals, and it should not be considered as an outright result of economic development. It is possible that gender equality may fall behind economic development, and may be even exposed to stagnation after a certain threshold level of income is achieved. This associates with the issue of gender equality decoupling, which is a subject for future research all by itself.

The 2007/2008 global financial crisis caused gender gap in employment to be higher than its pre-crisis level in the period 2009-2012 even in the most advanced economies (UN Women, 2014). The crisis endangered the gender equality process towards stagnation and even reversal. If the process may be exposed to devolution in the most advanced economies, then it may even experience reversals in developing ones during economic downturns. In this sense, Chapter 5 aims to determine two conditions over the period 1984-2014 for G7 countries that are the most advanced economies in terms of income and social transformation: (i) whether the impacts of shocks on the labor market gender equality are persistent or transitory in nature, when structural breaks are taken into consideration, and (ii) if and to what extent income affects the labor market gender equality, given structural breaks in the process. To this end, the first step is to identify the stationarity features of FLFPR and FSE, the two indicators of gender equality in the labor market by Narayan and Popp (NP) (2010) unit root test that allows for endogenously determined structural breaks at two time points. The NP unit root test findings ascertain that FLFPR and FSE are characterized by non-stationary behaviors for all G7 countries in the sample period, which imply that the impacts of shocks on the labor market gender equality are permanent. The subsequent step is to estimate long-run and short-run income elasticities of FLFPR and FSE by also regarding identified structural break dates of these indicators through autoregressive distributed lag (ARDL) bounds testing approach of cointegration. The ARDL estimates reveal that GDP per capita has a positive impact on FLFPR in four and on FSE in five out of G7 countries in the period 1984-2014. In all, Chapter 5 concludes that in case of the lack of intervention, any direct or indirect (via income) negative shock on FLFPR and FSE has potential to cause permanent damages on the improvement in the labor market gender equality process in even the most developed economies.

CHAPTER 2

GENDER EQUALITY IN THE LABOR MARKET AND INCOME CONVERGENCE: THEORY AND EVIDENCE

2.1 Introduction

Gender discrimination in the overall labor market has negative economic outcomes. It leads to the misallocation of skills and talents in labor force via the exclusion of women from productive work within their capacity. Discrimination lowers the average skill level, and thus, lowers average productivity. Hence, gender discrimination distorts the efficient allocation of labor and overall productivity in an economy. The inefficiency caused by gender discrimination is inevitably reflected in economic growth and development (Klasen, 1999; Esteve-Volart, 2004; Klasen and Lamanna, 2009; Ferrant, 2015). Making use of women's skill and improving their status to the same extent as that of men in the labor market contributes to macroeconomic stability and economic growth (Tzannatos, 1999; Stotsky, 2006; Morrison, Raju and Sinha, 2007; Löfström, 2009). However, the literature is lacking in theoretical and empirical evidence regarding the impact of the labor market gender equality on income convergence. The income convergence hypothesis that sustains Solow (1956) and Swan (1956) exogenous growth model conjectures that per capita growth of a country is inversely related to its initial level of income per capita due to diminishing marginal returns principle. The idea has been one of the focus points of the empirical growth literature for many years. Early studies in this direction are Abramovitz (1986), Baumol (1986), De Long (1988), Barro (1991), Barro and Sala-i Martin (1992), Levine and Renelt (1992) and Mankiw, Romer and Weil (1992). The cross-sectional regression analysis were later broadened by (i) panel data studies such as Islam (1995), Caselli, Esquivel and Lefort (1996), Evans and Karras (1996a, 1996b), Evans (1997), Lee, Pesaran and Smith (1997), Bond et al. (2001), Hoeffler (2002), and (ii) time series approaches such as Bernard and Durlauf (1995, 1996).

The conditional income convergence suggests that income per capita of a country converges to a long-run level, which is determined by the structural characteristics of that county. Therefore, the hypothesis implies that countries having similar structural characteristics converge to similar balanced growth paths in terms of per capita income in the long run. The model can be then considered in a broader perspective, for instance, in which a social characteristic of a country is included. This chapter investigates the role of gender equality in the labor market on conditional income convergence both theoretically and empirically at macro level. To this end, a gender-augmented convergence equation is developed in the framework of Solow-Swan neoclassical growth model, believed to be the first of its kind, which shows the positive role of the labor market gender equality on income convergence theoretically. This research takes an innovative approach, decomposing the total labor stock into female and male labor stocks in the production function. This novel modeling approach reveals a concave relationship between the labor market gender equality and income, and hence a 'golden rate' that maximizes income. In particular, when production elasticities of women and men workers are identical, it is shown that income is maximized when the share of female labor stock in total labor stock is 0.5, that is, the golden rate implies absolute gender equality in the labor market. The second innovative approach in this research is that the labor market gender equality is measured by the female share in employment, rather than the female labor force participation rate. The idea is that the latter may have drawbacks in the actual assessment of absolute gender equality in the labor market, since it does not distinguish between women who are employed and those actively searching for a job. On the other hand, the female share in employment counts in women of working age, who are engaged in economic activities to produce goods and/or to provide services, hence reflects the status of women in the labor market more accurately. This issue has not previously been discussed in any detail, and therefore it can be considered as a contribution to the literature.

In the tradition of Islam (1995), Caselli et al. (1996), Bond et al. (2001) and Hoeffler (2002), the empirical part of this chapter employs a dynamic panel data approach to estimate the convergence equation, using 5-year span panel data of 34 OECD countries in the period 1951-2010. Due to their continued labor market progress in increasing participation and employment of women over the post-war period, OECD countries are suitable candidates for the measurement of the role of gender equality on income convergence. Much of the economic growth in the OECD area is attributable to the increasing female employment in recent decades. In particular, the converging rates of female and male employment across the OECD has accounted for a quarter of annual economic growth since 1995 (OECD, 2008). In addition, the OECD launched its OECD Gender Initiative in 2010 with the aim of improving policies and initiatives for greater gender equality in employment and a more efficient utilization of everyone's skills as new sources of economic growth (OECD, 2011). Besides the general effect of increase in efficiency on income convergence due to a larger pool of workers to draw from, there are several channels through which greater gender equality in the labor market contributes to income convergence. The first channel is the structural changes in the economy. Across the OECD, the shift of employment from agriculture and manufacturing towards service sector throughout the 1980s and the 1990s has led to the greater participation of women in the labor market (OECD Employment Outlook, 2002). As the share of service sector in GDP increases, there is a greater demand for more flexible and low-cost labor, for which women are more suited. On average across the OECD, women are paid 16% less than men (OECD, 2011). Therefore, it is natural to expect that the female labor force will increase both proportionally and in absolute numbers in the service sector, rather than simply being a general substitute for male labor force. The second channel is that the labor supply behavior of women has changed via the change in family patterns, which stems from the importance of women's earnings in household income (OECD Employment Outlook, 2002). In addition, the change in social norms through rising levels of education, delayed marriage and lower fertility rates have led to the feminization of the labor market (Goldin, 1990; Ahn and Mira, 2002; Bloom et al., 2009). As a consequence of these two channels, the national income rises due to the increase in labor factor income. The final channel is the changing nature of competition due to globalization, which has reflected the commitment of governments to increase female employment rates. This has led to institutional change in the labor market: firms started to supply more part-time or irregular jobs with lower wages, which increases the female labor market participation (Standing, 1999; OECD Employment Outlook, 2002). Productive but underpaid female labor provides additional resources for investment through lowering unit labor costs, and increases competitiveness of firms and countries, which in turn stimulates productivity and economic growth (Ertürk and Çağatay, 1995; Seguino, 2000; Blecker and Seguino, 2002; Braunstein, 2012). Hence, the female labor market participation leads to a rise in income. In OECD countries, on average, 26% of females, and less than 7% of males are engaged in part-time employment (OECD Employment Outlook, 2002).

The empirical analysis of the role of the labor market gender equality on conditional income convergence is performed using System Generalized Method of Moments (GMM) estimator proposed by Arellano and Bover (1995) and Blundell and Bond (1998). This methodology is preferred, since it (i) fits for empirical growth models with a small number of time periods and a relatively large number of countries in nature, (ii) overcomes the issues of fixed effects and endogeneity of regressors, while avoiding dynamic panel bias, and (iii) yields more consistent and efficient parameter estimations as compared to other panel data estimators.[5] The regressions are based on 5-year span panel data to eliminate the cyclical component and to reduce serial correlation (Islam, 1995; Caselli et al., 1996). First, the textbook model is estimated. In the textbook model, the growth of GDP per worker depends on GDP per worker of the previous period, the saving rate and the effective depreciation rate of capital. Second, the textbook model is augmented with the female share in employment to analyze the impact of the labor market gender equality on income convergence. Finally, in order to increase the explanatory power of the female share in employment on income convergence, three macroeconomic control variables, including trade openness, foreign direct investment (FDI) inward stock and income inequality are added to the model.

[5] The way System GMM handles endogeneity problem is the key reason of choosing this method. The causality between income and gender equality in the labor market runs in both directions. There is also an extensive body of research regarding the (linear) impact of income (economic growth/development) on gender equality in the labor market, e.g., Weiss et al. (1976), Clark et al. (1991), Forsythe et al. (2000), Richards and Gelleny (2007), Duflo (2012).

The empirical findings of both the textbook model and the augmented model reveal a concrete evidence of GDP per worker convergence of OECD countries in the sample period. The female share in employment is positive and statistically significant in all estimations of the augmented model. In addition, all control variables are found to be statistically significant when included one by one in the augmented model, indicating that each of them affects the relationship between the female share in employment and income convergence. The implied speed of convergence, which is 3.1% per period in the textbook model, rises to 3.9% per period in the model augmented with the female share in employment. Moreover, the inclusion of the control variables in the augmented model enhances the implied convergence rate to 4.3%-7.8% per period. Hence, the strong empirical evidence suggests that the labor market gender equality has a considerable positive contribution on income convergence of OECD countries in the period 1951-2010, and also that this contribution becomes stronger when trade openness, FDI inward stock and income inequality are taken into account. All System GMM estimations are consistent and robust in terms of the validity of instrument set and of the expected signs of coefficients of the determinants of income convergence in both the textbook and the augmented model. All in all, this chapter testifies that a more gender-equal labor market is a driving force of income convergence of OECD countries. Hence, policy makers should develop incentive mechanisms and institutional adjustments in the labor market, which help remove gender-based barriers for female workers, and enhance female participation in income-generating activities.

The rest of the chapter is as follows: Section 2.2 develops a gender-augmented conditional income convergence equation. Section 2.3 is reserved for the empirical approach to the role of the labor market gender equality on conditional income convergence. Section 2.4 summarizes the chapter and provides some policy implications.

2.2 Theoretical Model

Suppose that production function is defined as:

$$Y = K^{\alpha} \cdot (A \cdot L_M)^{\beta} \cdot (A \cdot L_F)^{1-\alpha-\beta} \tag{2.1}$$

In Equation (2.1), K is physical capital, L_M is male labor force, L_F is female labor force, A is the overall technological progress, defined $A = A_0 \cdot e^{xt}$, and x is the exogenous rate of technological progress. Evidently, male and female labor forces are substitutes. The parameters $\alpha > 0$, $\beta > 0$, and $1 - \alpha - \beta > 0$ represent output elasticities (income shares) of capital, male labor, and female labor, respectively.[6] It is presumed that $L_M + L_F = L$.

The production function in effective per worker terms, $A \cdot L$, is expressed in Equation (2.2).

$$\tilde{y} = \tilde{k}^{\alpha} \cdot (MSE)^{\beta} \cdot (FSE)^{1-\alpha-\beta} \tag{2.2}$$

where $\tilde{y} = \frac{Y}{AL}$, $\tilde{k} = \frac{K}{AL}$, and $MSE \equiv \frac{L_M}{L}$ is the share of male labor force and $FSE \equiv \frac{L_F}{L}$ is the share of female labor force in the total labor stock. Hence, $MSE + FSE = 1$. It is assumed that male and female labor forces grow at the same rate, n, which implies that the total labor force grows at that particular rate. Hence, MSE and FSE are constant. Given that the ultimate aim is to develop a testable income convergence equation rather than to explain the sources of changes in the labor market gender equality, this is an innocent and acceptable assumption. The aspect that is novel and contributive to the literature of this approach is that it allows the exact measurement of the role of the labor market gender equality on income convergence. To understand this, the production function in effective per worker is assessed in natural logarithm in Equation (2.3).

$$\ln[\tilde{y}] = \alpha \cdot \ln[\tilde{k}] + \beta \cdot \ln[MSE] + (1 - \alpha - \beta) \cdot \ln[FSE] \tag{2.3}$$

In Equation (2.3), ln denotes natural logarithm. The first and the second derivatives of $\ln[\tilde{y}]$ with respect to FSE shows that there is a concave relationship between the two, and that $\ln[\tilde{y}]$ is maximized (for given capital per effective capita) when $FSE =$

[6] For the matter of theoretical perfectness, the production elasticities of male and female labor are denoted by different parameters. This does not, however, leave out the possibility of $\beta = 1 - \alpha - \beta$, which implies $\beta = \frac{1-\alpha}{2}$. Indeed, it is assumed that the latter is valid in the empirical analysis.

$\frac{1-\alpha-\beta}{1-\alpha}$.[7] It would not be inappropriate to describe this as the 'golden rate of FSE'. Interestingly, if male and female workers' production elasticities were equal, that is, $\beta = 1 - \alpha - \beta$, $\ln[\tilde{y}]$ would be maximized when $FSE = 0.5$, that is the golden rate of gender equality in the labor market would imply 'the absolute equality'. In conclusion, this theoretical approach incorporates an intuitive measure of the role of the labor market gender equality on income.

The second innovative approach is the measurement of the labor market gender equality by the share of females in total employment, although the literature rather utilizes the female labor force participation rate. The motivation is that using the latter may be misleading to reflect the absolute gender equality, since it does not regard the classification in the composition of female labor force in terms of employed versus actively looking for a job.

In a Solow (1956) framework, the fundamental equation of growth in effective per worker would then be:

$$\dot{\tilde{k}} = s \cdot \tilde{k}^{\alpha} \cdot (MSE)^{\beta} \cdot (FSE)^{1-\alpha-\beta} - (n + \delta + x) \cdot \tilde{k} \qquad (2.4)$$

In Equation (2.4), a dot over the effective capital per worker indicates its time derivative, namely, $\dot{\tilde{k}} = \frac{d\tilde{k}}{dt}$. s is the physical capital accumulation rate (equivalently saving or investment rate), and $n + \delta + x$ is the employed population growth rate (n) adjusted for the physical depreciation rate (δ) and technological progress rate (x), that is equivalent to the effective depreciation rate of capital.

The steady state value of effective capital per worker is then $\tilde{k}_{ss} = \left(\frac{s}{n+\delta+x}\right)^{\frac{1}{1-\alpha}} \cdot (MSE)^{\frac{\beta}{1-\alpha}} \cdot (FSE)^{\frac{1-\alpha-\beta}{1-\alpha}}$, which implies $\tilde{y}_{ss} = \left(\frac{s}{n+\delta+x}\right)^{\frac{\alpha}{1-\alpha}} \cdot (MSE)^{\frac{\beta}{1-\alpha}} \cdot (FSE)^{\frac{1-\alpha-\beta}{1-\alpha}}$.

[7] The first and the second derivatives of $\ln[\tilde{y}]$ with respect to FSE are $\frac{-\beta}{(1-FSE)} + \frac{1-\alpha-\beta}{FSE}$ and $-\left[\frac{\beta}{(1-FSE)^2} + \frac{1-\alpha-\beta}{FSE^2}\right]$, respectively. It is worthy of note that the same result could have been obtained from the first and the second derivation of $\tilde{y}$, with respect to FSE in Equation (2.2).

In order to determine the income convergence implication of the labor market gender equality, the log differentiation of production function in per capita terms is obtained at first in Equation (2.5).

$$\hat{\tilde{y}} = \alpha \cdot \hat{\tilde{k}} \tag{2.5}$$

where $\hat{\tilde{y}} = \frac{\dot{\tilde{y}}}{\tilde{y}}$ and $\hat{\tilde{k}} = \frac{\dot{\tilde{k}}}{\tilde{k}}$. Then, the fundamental equation of growth in terms of $\tilde{y}$ can be expressed. To this end, Equation 2.4 is divided by $\tilde{k}$, and then $\hat{\tilde{k}}$ and $\tilde{k}$ are expressed in $\hat{\tilde{y}}$ and $\tilde{y}$, respectively, which yields Equation (2.6).[8]

$$\frac{d\ln[\tilde{y}]}{dt} = \alpha \cdot \left[s \cdot e^{\left(\frac{\alpha-1}{\alpha}\right)\ln[\tilde{y}]} \cdot (MSE)^{\frac{\beta}{\alpha}} \cdot (FSE)^{\frac{1-\alpha-\beta}{\alpha}} - (n+\delta+x)\right] \equiv \phi(\ln[\tilde{y}]) \tag{2.6}$$

where $\phi(\ln[\tilde{y}])$ underlines the fact that the RHS of (2.6) is a function of $\ln[\tilde{y}]$. Through Taylor expansion, (2.6) can be log-linearized:

$$\frac{d\ln[\tilde{y}]}{dt} \approx \phi\ln[\tilde{y}_{ss}] + \phi'(\ln[\tilde{y}_{ss}])[\ln[\tilde{y}] - \ln[\tilde{y}_{ss}]] \tag{2.7}$$

In Equation (2.7), $\phi\ln[\tilde{y}_{ss}]$ is steady-state value of $\phi(\cdot)$, and $\phi'(\ln[\tilde{y}_{ss}])$ is the derivative of $\phi(\cdot)$ evaluated at steady state. It is well known that $\phi\ln[\tilde{y}_{ss}] = 0$ in a Solovian setup without (exogenous) technological change. Through the standard convergence algebra, it is easy to get the following gender-augmented income convergence equation:

$$\ln[y(t_2)] - \ln[y(t_1)] = -(1-e^{-\nu\cdot\tau}) \cdot \ln[y(t_1)] + (1-e^{-\nu\cdot\tau}) \cdot \frac{\alpha}{1-\alpha} \cdot \ln[s] - (1-e^{-\nu\cdot\tau}) \cdot \frac{\alpha}{1-\alpha} \cdot \ln[n+\delta+x] + (1-e^{-\nu\cdot\tau}) \cdot \frac{\beta}{1-\alpha} \cdot \ln[MSE] + (1-e^{-\nu\cdot\tau}) \cdot \frac{1-\alpha-\beta}{1-\alpha} \cdot \ln[FSE] + (1-e^{-\nu\cdot\tau}) \cdot \ln[A_0] + x(t_2 - e^{-\nu\cdot\tau}t_1) \tag{2.8}$$

In Equation (2.8), y is income per worker, $\nu = (1-\alpha) \cdot (n+\delta+x)$ is the speed of convergence to steady state, t_2 and t_1 such that $t_2 > t_1$ are two points in time, and

[8] Please note that $\tilde{y} = \tilde{k}^{\alpha} \cdot (MSE)^{\beta} \cdot (FSE)^{1-\alpha-\beta}$ implies $\tilde{k} = \tilde{y}^{\frac{1}{\alpha}} \cdot (MSE)^{-\frac{\beta}{\alpha}} \cdot (FSE)^{-\frac{1-\alpha-\beta}{\alpha}}$.

$\tau = t_2 - t_1$. Two further assumptions are in need before embarking on to empirical analysis. First, it is more realistic to use the absolute gender equality version of Equation (2.8) in the empirical part of the chapter. Second, as a final adjustment, MSE should be dropped from Equation (2.8) by using $MSE + FSE = 1$, in order to avoid a multicollinearity problem. Under these assumptions, gender-augmented income convergence equation becomes:

$$\ln[y(t_2)] - \ln[y(t_1)] = -(1 - e^{-\nu\cdot\tau}) \cdot \ln[y(t_1)] + (1 - e^{-\nu\cdot\tau}) \cdot \frac{\alpha}{1-\alpha} \cdot \ln[s] - (1 - e^{-\nu\cdot\tau}) \cdot \frac{\alpha}{1-\alpha} \cdot \ln[n + \delta + x] + \frac{(1-e^{-\nu\cdot\tau})}{2} \cdot \ln[FSE] + (1 - e^{-\nu\cdot\tau}) \cdot \ln[A_0] + x(t_2 - e^{-\nu\cdot\tau} t_1) \quad (2.9)$$

Equation (2.9) is adopted for the empirical estimation of the role of the labor market gender equality on income convergence.

2.3 Empirical Analysis

2.3.1 Methodology

In the empirical analysis, System GMM estimator proposed by Arellano and Bover (1995) and Blundell and Bond (1998) is employed. There are five major motivations to utilize the System GMM within the framework of panel data methodology. Firstly, though widely used in the convergence literature, Ordinary Least Squares (OLS) levels and Within Groups estimators lead to biased and inconsistent estimations in dynamic panel data framework. OLS levels and Within Groups estimations are biased and inconsistent, since (i) OLS levels omits unobserved time invariant country effects, and (ii) Within Groups takes account of the unobserved country-specific effects with a fixed time period in dynamic panel data model (Hsiao, 2014; Nickell, 1981). The coefficients of the lagged dependent variable estimated by OLS levels and Within Groups are regarded as approximate upper and lower bounds, respectively (Bond et al., 2001; Hoeffler, 2002). The System GMM would be rather employed as a robust technique, since it yields unbiased estimations (Arellano and Bover, 1995; Blundell and Bond, 1998, 2000; Blundell, Bond and

Windmeijer, 2001). Secondly, the System GMM provides consistent and efficient parameter estimations in a regression, in which independent variables are not strictly exogenous, i.e., they are correlated with past and current realizations of the error term, and/or in which heteroscedasticity and autocorrelation within individuals exist (Roodman, 2009a). This estimator gets over the endogeneity issue by instrumenting the lagged dependent variable and/or any other endogenous variables with variables, which are thought to be uncorrelated with the fixed effects (Nickell, 1981; Roodman, 2009a). Thirdly, the System GMM is more efficient than the Difference GMM estimator proposed by Arellano and Bond (1991), since it enables the inclusion of more instruments with an additional assumption that the first differences of instruments are uncorrelated with the fixed effects (Roodman, 2009a). In addition, the System GMM yields efficient estimations in cases where the series are close to being random walks, whereas the Difference GMM estimations can be exposed to large finite sample biases in these cases (Blundell and Bond, 1998). The coefficient of the lagged dependent variable estimated by the Difference GMM tends to be biased downwards, like the Within Groups estimation, if the instruments are weak, and the number of time series observations is small (Blundell and Bond, 2000; Hoeffler, 2002). Fourthly, the System GMM estimator is designed for panel data sets, which consist of small time dimension, and relatively large cross-sectional dimension (Blundell and Bond, 1998, 2000; Blundell et al., 2001). Finally, this estimator is highly suggested for empirical growth models due to its fit for linear equations with one dynamic dependent variable, additional explanatory variables and fixed effects (Bond et al., 2001; Roodman, 2009b).

The first-order autoregressive panel data model is given by:

$$y_{it} = \theta \cdot y_{i,t-1} + \gamma \cdot X_{it} + \mu_i + \phi_t + \varepsilon_{it} \; for \; i = 1, \dots, N \; and \; t = 2, \dots, T$$
$$u_{it} = \mu_i + \phi_t + \varepsilon_{it} \qquad (2.10)$$

In Equation (2.10), $y_{i,t-1}$ is the lagged dependent variable, and X_{it} is the vector of explanatory variables. μ_i and ϕ_t measure country-specific (fixed) effects and time-specific intercepts, respectively. ε_{it} denotes the idiosyncratic shocks, which vary across countries i, and time periods t. u_{it} is the error term. The first step of the estimation procedure is the transformation of the variables (by first-differencing) to eliminate μ_i from u_{it}, hence to overcome the issue of endogeneity arising from the correlation between $y_{i,t-1}$ and u_{it}:

$$\Delta y_{it} = \theta \cdot \Delta y_{i,t-1} + \gamma \cdot \Delta X_{it} + \Delta\phi_t + \Delta\varepsilon_{it}\ for\ i = 1, \ldots, N\ and\ t = 3, \ldots, T \qquad (2.11)$$

It is assumed that the transitory errors are independent across countries and serially uncorrelated:

$$E(\varepsilon_{it}\varepsilon_{is}) = 0\ for\ t \neq s \qquad (2.12)$$

and that the initial conditions satisfy:

$$E(y_{i1}\varepsilon_{it}) = 0\ for\ t \geq 2 \qquad (2.13)$$

These assumptions imply that $y_{i,t-1}$ is predetermined with respect to ε_{it}.

Although μ_i are removed from the regression, $y_{i,t-1}$ is still potentially endogenous, since $y_{i,t-1}$ in $\Delta y_{i,t-1}$ is correlated with $\varepsilon_{i,t-1}$ in $\Delta\varepsilon_{it}$. In order to get over this issue, Arellano and Bond (1991) propose the moment restrictions given in Equations (2.14) and (2.15):

$$E\left(y_{i,t-s}\Delta\varepsilon_{it}\right) = 0\ for\ t = 3, \ldots, T\ and\ s \geq 2 \qquad (2.14)$$

$$E\left(X_{i,t-s}\Delta\varepsilon_{it}\right) = 0\ for\ t = 3, \ldots, T\ and\ s \geq 2 \qquad (2.15)$$

Nevertheless, Blundell and Bond (1998) show that the lagged levels of the explanatory variables are weak instruments for the first-differenced equation, when these variables are persistent over time. In this case, the estimations are exposed to finite sample bias due to the weak correlation between the instruments and the endogenous variables, especially in small samples. The solution by Blundell and Bond (1998) is to compose a system with two sets of equations. One set of equations is the transformed equation in first differences, for which the lagged levels of y_{it} and

X_{it} are used as instruments. The other set of equations is the original equation in levels, for which the lagged first differences of y_{it} and X_{it} are used as instruments. The additional moment restrictions developed by Blundell and Bond (1998) are in Equations (2.16) and (2.17):

$$E\left[\Delta y_{i,t-1}(\mu_i + \varepsilon_{it})\right] = 0 \; for \; t = 3, \ldots, T \tag{2.16}$$

$$E\left[\Delta X_{i,t-1}(\mu_i + \varepsilon_{it})\right] = 0 \; for \; t = 3, \ldots, T \tag{2.17}$$

For the consistency of the System GMM estimations, it is important to ensure four key conditions:

i. There should be no serial correlation in the error term. Arellano-Bond (1991) test identifies the first and the second order serial correlations in the first-differenced residuals. The second-order correlation in first differences is taken into account to analyze the first-order serial correlation in levels, since this will determine the correlation between $\varepsilon_{i,t-1}$ in $\Delta\varepsilon_{i,t-1}$ and $\varepsilon_{i,t-2}$ in $\Delta\varepsilon_{i,t-2}$ (Roodman, 2009a). Arellano-Bond test for AR(2) in first differences reports the p-values for the null hypothesis of no second-order serial correlation in the first-differenced residuals.

ii. The instruments should not be correlated with the error term to ensure the instrument validity. Hansen (1982) test of over-identifying restrictions reports the p-values for the null hypothesis of instrument validity.

iii. The additional moment restrictions by Blundell and Bond (1998) should be valid. The Difference-in-Hansen test reports the p-values for the null hypothesis of the validity of additional moment conditions.

iv. As a rule of thumb, the number of instruments should be smaller than or equal to the number of groups in a regression to avoid finite sample bias caused by overfitting (Roodman, 2009b).

The following dynamic panel data equation is estimated in order to measure the role of the female share in employment on income convergence:

$$\ln[y_{it}] = \theta \cdot \ln\left[y_{i,t-1}\right] + \gamma_1 \cdot \ln[s_{it}] + \gamma_2 \cdot \ln[n_{it} + \delta + x] + \gamma_3 \cdot \ln[FSE_{it}] + \gamma_4 \cdot \ln[Z_{it}] + \mu_i + \phi_t + \varepsilon_{it} \tag{2.18}$$

The LHS of Equation (2.18), $\ln[y_{it}]$ denotes income per worker over a 5-year time period. The determinants of income convergence take place on the RHS. θ is the coefficient of income per worker of the previous 5-year time period, $\ln[y_{i,t-1}]$. This coefficient is expected to be between 0 and 1. Accordingly, $\theta^* = \theta - 1$ is expected to be between -1 and 0, which is consistent with the convergence hypothesis.[9] The coefficient γ_1 shows the contribution of the saving rate, $\ln[s_{it}]$, γ_2 measures the impact of effective depreciation rate of capital, $\ln[n_{it} + \delta + x]$, γ_3 measures the contribution of the female share in employment, $\ln[FSE_{it}]$, on income convergence in a 5-year time span. In addition, γ_4 denote the respective coefficients of the vector of control variables in a 5-year time span, $\ln[Z_{it}]$, which are used to increase the explanatory power of female share in employment on income convergence. These variables include trade openness, FDI inward stock and income inequality, which are expected to affect the relationship between the female share in employment and income convergence due to the following motivations:

i. Trade openness: The literature argues that the enlargement in trade due to globalization provides job opportunities for women, and promotes the participation of them in export-oriented industries (Wood, 1991; Çağatay and Özler, 1995; Standing, 1999; Kucera and Milberg, 2000; Bussmann, 2009).
ii. FDI inward stock: The literature asserts that female employment in export-oriented industries and manufacturing sector is high in countries with sizeable FDI inflows (Joekes and Weston, 1994; Braunstein, 2002). FDI provides better economic rights for women (Braunstein and Brenner, 2007).
iii. Income inequality: In the literature, there is evidence that females are less likely to be excluded from labor force in lower income inequality societies (Semyonov, 1980).

Finally, time dummies are included in regressions, since this will make the assumption of no correlation across individuals in the idiosyncratic disturbances more likely to hold (Roodman, 2009a).

[9] Please note that θ^* would be the coefficient of $\ln[y_{i,t-1}]$ if the dependent variable were in growth form, that is, $\ln[y_{it}] - \ln[y_{i,t-1}]$.

2.3.2 Data

The empirical analysis is based on unbalanced panel data of 34 OECD countries over the period 1951-2010.[10] In Solow model regressions, it is more advisable to use per worker variables, since the model is based on a production function and not every person is included in production activities (Hoeffler, 2002). In the tradition of Mankiw et al. (1992), per worker GDP and the growth rate of employment is used in the analysis.[11] The data of real GDP at 2005 constant prices in millions US$, the employed population in millions and the share of gross capital formation in GDP (equivalent to the saving rate) are obtained from Penn World Table 8.0 of Feenstra, Inklaar and Timmer (2015). The physical depreciation rate and the technological progress rate are assumed to be time and country invariant variables, and the sum of them is regarded as 5%, as such in Mankiw et al. (1992), Islam (1995) and Caselli et al. (1996). The data of female employment for the age range of 15-64 in thousands are retrieved from OECD Labour Force Statistics and ILOSTAT- International Labour Organization (ILO) Database of Labour Statistics. The source of data of trade openness at 2005 constant prices in percentage is Penn World Table 7.1 of Heston, Summers and Aten (2012). The data of the share of FDI inward stock in GDP are extracted from United Nations Conference on Trade and Development (UNCTAD) Statistics. The income inequality is measured by Gini index. The data of Gini coefficients of income (disposable income, post taxes and transfers) are from OECD Income Distribution Database (IDD). Following Islam (1995) and Caselli et al. (1996), 5-year averages are calculated to reduce serial correlation problem and to avoid the effect of business cycle fluctuations. Hence, 12 data (time) points for each of 34 countries are obtained, e.g., 1955, 1960, ..., 2010. Table 2.1 presents the summary statistics of the series in panel data set.

[10] OECD countries are Australia, Austria, Belgium, Canada, Chile, Czech Republic, Denmark, Estonia, Finland, France, Germany, Greece, Hungary, Iceland, Ireland, Israel, Italy, Japan, Republic of Korea, Luxembourg, Mexico, the Netherlands, New Zealand, Norway, Poland, Portugal, Slovak Republic, Slovenia, Spain, Sweden, Switzerland, Turkey, United Kingdom and the United States of America.

[11] Islam (1995) and Caselli et al. (1996) use per capita variables. Hoeffler (2002) runs all regressions in both per capita and per worker terms, but yields similar results.

TABLE 2.1: Descriptive Statistics of 5-year span data, 1951-2010

Variables	Observations	Mean	Std. Dev.	Min.	Max.
GDP (millions)	368	621401.3	1412449	1096.86	12964400
Employed population (millions)	364	13.26	21.46	0.07	144.71
Gross capital formation (% of GDP)	374	24.98	5.98	10.33	46.78
Effective depreciation rate of capital (%)	362	6.06	1.47	-0.32	10.91
Female share in employment (%)	284	39.41	6.64	19.77	50.54
Trade openness (%)	371	51.47	43.07	3.86	313.19
FDI inward stock (% of GDP)	212	26.99	34.19	1.00	249.71
Income inequality (%)	138	30.73	6.57	19.80	51.90

Note: Std. Dev., Min. and Max. denote standard deviation, minimum and maximum, respectively.

2.3.3 Findings

Table 2.2 presents the panel regression results from estimating Equation (2.18) by one-step System GMM estimator with 5-year span data of 34 OECD countries over the period 1951-2010.[12] For System GMM, the two-step estimator is more efficient than the one-step estimator. Nevertheless, Monte Carlo studies indicate that the efficiency gain is small, and that the two-step estimator converges only slowly to its asymptotic distribution. The asymptotic standard errors related to the two-step GMM estimators can be seriously biased downwards in finite samples (Blundell and Bond, 1998; Hoeffler, 2002). Hence, following Hoeffler (2002), one-step System GMM estimations are reported. In Table 2.2, column (1) shows the estimations of the textbook model, in which GDP per worker of the previous 5-year time period (lagged dependent variable), the saving rate and the effective depreciation rate of capital are the determinants of GDP per worker convergence; column (2) presents the estimations of the augmented model, which additionally assesses the impact of female share in employment on GDP per worker convergence; columns (3-6) present the estimations of the augmented model, respectively including trade openness, FDI inward stock, income inequality and all three control variables.

In Table 2.2, the lagged dependent variable is assumed to be predetermined, the saving rate and the effective depreciation rate of capital, respectively, are regarded as

[12] The command “xtabond2” is used in Stata (v.13) for System GMM estimations, and the instrument matrix is collapsed with the command “collapse” available in Stata, as mentioned in Roodman (2009b).

endogenous and exogenous regressors for all regressions.[13] The female share in employment and the control variables take part in the augmented model as endogenous and exogenous regressors, respectively.[14] In addition, time dummies for 12 data points are included in all regressions; however, their coefficients are not reported in order to conserve space. The first row shows $\hat{\theta}$, the estimated coefficient of the lagged dependent variable, which is expected to be between 0 and 1. This implies $\widehat{\theta^*} = \hat{\theta} - 1$ to be between -1 and 0, which is an evidence of GDP per worker convergence. The implied speed of convergence is determined from the theoretical model such that $\nu = -\frac{\ln \hat{\theta}}{\tau}$, where $\tau = t_2 - t_1 = 5$. Hence, a lower $\hat{\theta}$, or a higher $\widehat{\theta^*}$ (in absolute value) indicates a higher implied speed of convergence in the model. The implied speed of convergence in the augmented model is expected be higher than that in the textbook model. In addition, the inclusion of the control variables in the augmented model is expected to enhance the implied convergence rate.

The coefficient of lagged GDP per worker is between 0 and 1, and statistically significant at 1% level in all estimations, which provides a strong evidence of GDP per worker convergence of OECD countries in sample period. The estimations of the other determinants of income convergence in the neoclassical model, namely the saving rate and the effective depreciation rate of capital, are statistically significant and consistent with the theoretical expectation (columns (1)-(6) of Table 2.2). In accordance with a priori expectations, the textbook model yields the lowest speed of convergence, which is 3.1 % per period (column (1) of Table 2.2). It rises to 3.9% per period when the model is augmented with the female share in employment (column (2) of Table 2.2). Moreover, the female share in employment is positive and statistically significant at either 1% or 5% level in all estimations of the augmented model, indicating that it contributes positively to GDP per worker convergence (columns (2)-(6) of Table 2.2).

[13] In empirical growth models, the saving (investment) rate is assumed to be endogenous, since there may be correlation between current investment and current shocks to income, as well as feedback from past shocks to income. On the other hand, the current effective depreciation rate of capital is assumed to be exogenous, in the sense of being uncorrelated with shocks to income in both the current and the previous 5-year time periods (Hoeffler, 2002).

[14] The endogeneity issue between income and the labor market gender equality is treated correctly in the regressions by System GMM estimator.

TABLE 2.2: System GMM Estimations of Income Convergence from a panel of 5-year span data, 1951-2010

Variables and Statistics	Dependent variable: ln[GDP per worker]					
	(1)	(2)	(3)	(4)	(5)	(6)
ln[GDP per worker(-1)]	0.856***	0.822***	0.772***	0.789***	0.806***	0.678***
	(0.02)	(0.03)	(0.03)	(0.07)	(0.03)	(0.09)
ln[saving rate]	0.291***	0.305***	0.270***	0.220***	0.044***	0.068**
	(0.05)	(0.08)	(0.08)	(0.07)	(0.01)	(0.03)
ln[effective depreciation rate of capital]	-0.082***	-0.057*	-0.091***	-0.052**	-0.066*	-0.217***
	(0.02)	(0.03)	(0.03)	(0.03)	(0.03)	(0.08)
ln[female share in employment]	-	0.089**	0.264**	0.216**	0.262***	0.190***
		(0.04)	(0.13)	(0.10)	(0.10)	(0.07)
ln[trade openness]	-	-	0.042*	-	-	0.045
			(0.03)			(0.12)
ln[FDI inward stock]	-	-	-	0.016**	-	0.326
				(0.01)		(0.21)
ln[income inequality]	-	-	-	-	-0.118**	-0.105***
					(0.05)	(0.03)
Constant	0.850***	0.447	0.690*	0.359	1.446**	0.768***
	(0.19)	(0.46)	(0.41)	(0.43)	(0.60)	(0.19)
Implied ν	0.031	0.039	0.052	0.047	0.043	0.078
Instruments	18	22	24	30	23	34
Groups	34	34	34	34	34	34
Hansen test	0.39	0.12	0.11	0.18	0.33	0.63
Difference-in-Hansen test	0.76	0.69	0.87	0.72	0.55	0.33
AR(2)	0.25	0.13	0.25	0.16	0.19	0.40

Notes: "(-1)" denotes one lag of the corresponding variable. Heteroscedasticity-consistent standard errors are in parentheses. Windmeijer (2005) finite sample correction for standard errors is employed. The superscripts ***, ** and * denote the statistical significance at 1%, 5% and 10% levels, respectively.

When trade openness, FDI inward stock and income inequality are added one by one to the augmented model, the implied speed of convergence, respectively, increases to 5.2%, 4.7%, and 4.3%. Moreover, the estimations of trade openness, FDI inward stock and income inequality are statistically significant, indicating that each of them increases the explanatory power of the female share in employment on income convergence (columns (3)-(5) of Table 2.2). The implied speed of convergence is the highest in the augmented model including all three control variables, which is 7.8%, although trade openness and FDI inward stock are not statistically significant (column (6) of Table 2.2). The reason is that the inclusion of a regressor, though insignificant, strengthens the instrument set significantly in GMM regressions (Hoeffler, 2002). These findings reveal that the female share in employment has a positive and statistically significant contribution on income convergence of OECD countries in period 1951-2010. In addition, this contribution improves when openness, FDI inward stock and income inequality are included in the model.

As an empirical matter, four key conditions should be considered for the consistency check of the System GMM estimations. Firstly, the p-values of AR(2) statistics indicate that there is no second-order serial correlation in the first-differenced residuals in any of the estimations. Secondly, the p-values of Hansen test of over-identifying restrictions imply that the instrumental variables are not correlated with the error term in any of the estimations. Thirdly, the p-values of Difference-in-Hansen test show that additional moment restrictions of the System GMM estimations are valid. Finally, the rule of thumb is satisfied, since the number of groups is larger than or equal to the number of instruments in all estimations. It all comes to this that the overall performance of both the textbook model and the augmented model estimations is consistent and robust, since the instrument set used is valid, and the coefficients of lagged GDP per worker, the female share in employment and the control variables are consistent with a priori expectations.

2.4 Concluding Remarks and Policy Implications

This chapter examines the role of gender equality in the labor market on conditional income convergence both theoretically and empirically. In the theoretical part, a gender-augmented income convergence equation is developed in the framework of Solow-Swan neoclassical growth model. The contribution of the model to the literature is twofold. First, it ascertains a ‘golden rate’ that would imply the absolute labor market gender equality, and that maximizes income per worker. Second, it suggests the female share in employment as a new measure of the labor market gender equality. The model shows the positive role of the female share in employment on income convergence. In the empirical part, a dynamic panel data equation is utilized to test the impact of female share in employment on income convergence. The System GMM approach is applied to 5-year span panel data of 34 OECD counties over the period 1951-2010. The consistent and robust System GMM estimations for both the textbook and the augmented models yield concrete evidence of the conditional GDP per worker convergence of OECD countries in the sample period. In the estimations of the augmented model, the female share in employment is positive, which confirms the theoretical expectation. Furthermore, the augmented model yields a higher speed of convergence than the textbook model giving rise to the argument that female share in employment is greatly overlooked predictor of conditional income convergence. The estimations also show that this contribution is reinforced when trade openness, FDI inward stock and income inequality are taken into account in the model.

This chapter provides a strong evidence of the positive role of a more gender-equal labor market on income convergence. This finding has very important policy implications at two levels. First, whoever is responsible for economic policies of countries, whether OECD or non-OECD, should aim to achieve absolute gender equality in the labor market for the sake of higher income convergence. In practical terms, this entails the policies and institutional adjustments to enhance the female share in employment, since females are the disadvantageous group in most economies in the world. Ensuring gender-equal labor market opportunities may require legislative, social, political, and other acts. Second, policy makers should develop policies for protecting previous achievements, as there may be a tendency for some policy makers to return to pro-male policies, especially during extraordinary times, e.g., global and regional economic crises, political instability, and disturbances in variables related to the labor market. To this end, policy makers should constantly monitor and support gender equality in the labor market to preserve efficiency, which is measured in terms of income convergence. That is, the sustainability of these measures is at least as important as their initial achievement.

CHAPTER 3

THE 'GRAND GENDER CONVERGENCE' IN OECD LABOR MARKET

3.1 Introduction

The neoclassical income convergence hypothesis supposes that a country that is further from its own steady-state income level will grow faster than one closer to its own steady-state income level. Since the pioneering contributions of Abramovitz (1986), Baumol (1986) and De Long (1988), several studies, including Barro (1991), Barro and Sala-i Martin (1992), Mankiw et al. (1992), Islam (1995), Quah (1996, 1997), Caselli et al. (1996), Evans and Karras (1996a, 1996b), Evans (1997), Bond et al. (2001), Hoeffler (2002) and Mathunjwa and Temple (2007) fully establish this research field. In view of the fact that countries with similar structural characteristics will eventually converge to similar balanced growth paths in terms of income per capita, one interesting research focus is to verify whether a similar convergence behavior exists in social development indicators. This is an important question because utility is not only derived from the use of income, but also from social development. Intuitively, any social indicator in which income per capita is cause as much as effect, may mimic income convergence behavior because it is highly possible that convergence in one feeds back into the other. The literature clearly reflects the increasing interest in convergence in social indicators. For example, Neumayer (2003) shows strong cross-country evidence of convergence in life expectancy, infant survival, educational enrolment, and literacy. Hobijn and Franses (2001) study the existence and the extent of convergence in two social indicators, infant mortality rates and life expectancy at birth, but fail to show that convergence in GDP per capita implies convergence in these social indicators. Dorius (2008) examines global demographic convergence over the last 50 years, and finds that lagging countries are catching up with countries that had begun the transition to low fertility earlier. Finally, Royuela and García (2015) analyze economic and social convergence in Colombia at regional level using life expectancy, infant mortality, educational enrolment, and crime as social indicators, and find convergence in key social variables, but not in GDP per capita.

Goldin (2014:1091) observes that:

> Of the many advances in society and the economy in the last century, the converging roles of men and women are among the grandest. A narrowing has occurred between men and women in labor force participation, paid hours of work, hours of work at home, life-time labor force experience, occupations, college majors, and education, where there has been an overtaking by females.

Inspired by the literature on convergence in social indicators in general and the observation by Goldin (2014) in particular, this chapter aims to verify the convergence behavior of gender equality in the labor market empirically.

The literature on income and gender equality in the labor market indicates a mutual interaction between the two, which strengthens the conjecture that gender equality in the labor market may follow income convergence. Firstly, gender equality in the labor market has (positive) effect on income dynamics. There is a substantial body of literature showing that (i) improving the status of women, and hence, reducing gender inequality in the labor market contributes to economic growth and macroeconomic stability (Tzannatos, 1999; Stotsky, 2006; Morrison et al., 2007; Löfström, 2009), (ii) gender discrimination in the labor market hinders overall productivity and economic outcomes in the long run by obstructing efficient workforce allocation (Klasen, 1999; Esteve-Volart, 2004; Klasen and Lamanna, 2009; Ferrant, 2015). Secondly, income dynamics has (positive) impact on gender equality in the labor market, linear or nonlinear. Some studies conclude that gender equality increases linearly in conjunction with increasing levels of income, e.g., Weiss et al., 1976; Clark, 1991; Clark et al., 1991; Charles, 1992; Forsythe et al., 2000; Gray, Kittilson and Sandholtz, 2006; Richards and Gelleny, 2007; Duflo, 2012. However, others claim that while gender equality decreases in the initial stages of economic development, it begins to increase after a particular threshold of income level is achieved (Boserup, 1970; Pampel and Tanaka, 1986; Psacharopoulos and Tzannatos, 1989; Goldin, 1995; Tam, 2011).

This chapter is based on empirical evidence for 34 OECD countries in the period 1971-2010. The motivation for analyzing gender equality convergence in labor market of OECD countries is the ever-narrowing gender gap across the OECD area in the last three or four decades. This phenomenon is clearly due to the increasing female labor force participation rates, rather than to variations in the incidence of unemployment (OECD Employment Outlook, 2002). On average, the female labor force participation rate increased about 29% across OECD countries in the period 1970-2010 (OECD Labour Force Statistics). There are three main factors giving rise to the increasing female labor market participation, including socio-cultural, economic and institutional changes. Firstly, social norms, family patterns and household formation have changed. The growing desire of women for paid employment and economic independence, and increasing role of women's earnings in household income have caused the feminization of labor force (OECD Employment Outlook, 2002). Moreover, increasing educational levels, delayed marriage and lower fertility rates have increased the supply of women workers (Goldin, 1990; Ahn and Mira, 2002; Bloom et al., 2009). Secondly, the composition of income has changed structurally. The shift of employment from agriculture and manufacturing towards service sector has specifically favored the employment of women. Women have generally benefited from the increased availability of better quality employment (Mehra and Gammage, 1999). Across the OECD, on average, almost one in three women are employed in the service sector, particularly in sales, health and community services, and education (OECD Employment Outlook, 2002). Finally, institutional changes favoring part-time or irregular employment with lower wages due to competition, and those providing more flexible working-time arrangements have led to the increase in female employment rates (Standing, 1999; OECD Employment Outlook, 2002). In the OECD area, almost three out of four part-time jobs are held by women workers, and more than one in four women prefer to work part-time (OECD, 2008).

The female labor force participation rate (FLFPR) is a potential indicator of women's status and gender equality, since it reflects the social and economic empowerment of the labor market (Eastin and Prakash, 2013). Since the pioneering studies of Mincer (1962) and Cain (1966), studies using the female labor force participation as an indicator of women's labor market activity include Semyonov (1980), Hill (1983), Psacharopoulos and Tzannatos (1989), Clark et al. (1991), Bloom et al. (2009) and Mishra and Smyth (2010). However, FLFPR may be inaccurate to capture women's contribution to overall production efforts, since it also takes into account unemployed women who are actively looking for a job. Hence, in order to prevent a statistical artifact arising from measurement variation, and to enhance the reliability of the analysis, the female share in employment (FSE) is used as a second indicator of the labor market gender equality. The motivation to use FSE is that it directly reflects the extent to which women occupy production activities and economic positions in the labor market.

A formal approach to investigating the convergence process is to view "convergence picture" by plotting annualized (or averaged over a period) growth rates against initial levels of the variable of interest.[15] The graphical demonstration in Figure 3.1 and Figure 3.2, each plotting the average annual growth rate of a labor market gender equality indicator against its initial value, also supports the hypothesis 'gender equality convergence in the labor market' for OECD countries. In particular, Figure 3.1 presents a scatter plot of the average annual FLFPR growth rate for the period 1971-2010 against FLFPR in 1971 (or the earliest observation) for each OECD countries.[16] The plot indicates an inverse relationship between the average growth rates and the initial values of FLFPR, which suggests the existence of gender equality convergence in labor market of OECD counties.

[15] Romer (1987) is the first to present the "convergence picture" for unconditional income convergence. Then, Mankiw et al. (1992) demonstrate income convergence in unconditional sense and conditional on population growth and accumulation of physical and human capital.

[16] The initial year concerned is not 1971 for all countries due to the limitations in the availability of data for the labor market gender equality indicators. For example, if the initial observation year is 1980 for a country, then the plot shows the average annual FLFPR growth rate for the period 1980-2010 against FLFPR in 1980 for that country.

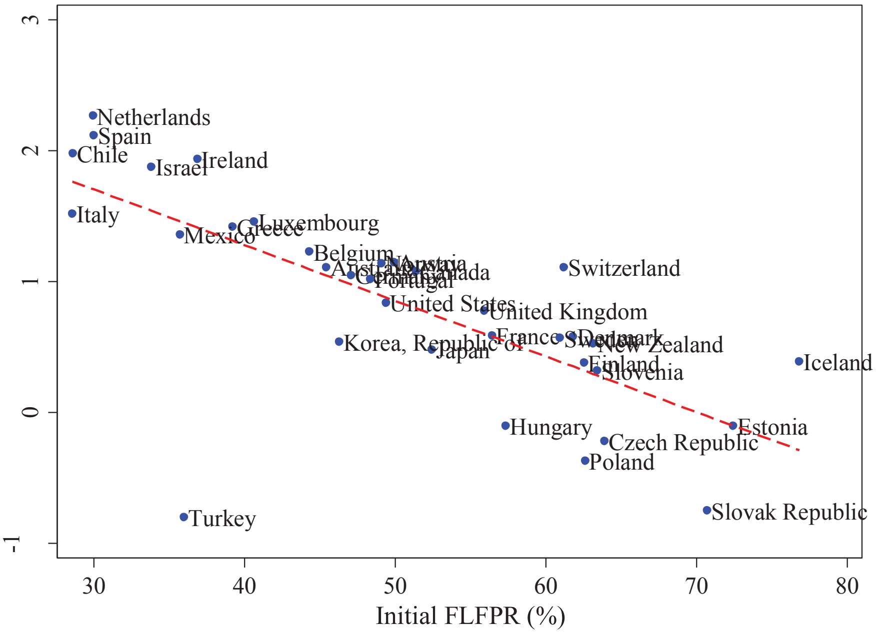

FIGURE 3.1: FLFPR Convergence of OECD countries, 1971-2010
Data Source: OECD Labour Force Statistics, ILOSTAT- ILO Database of Labour Statistics, Author's calculations.

In a similar manner, Figure 3.2 plots the average annual FSE growth rate from 1971 to 2010 against FSE in 1971 (or the earliest observation) for each OECD countries. The plot shows that FSE average growth rates are higher (lower) in the countries, for which the initial values of FSE are lower (higher), also supporting the argument that OECD countries have a strong tendency to converge in gender equality in the labor market. Figure 3.1 and Figure 3.2 indicate that Turkey is an outlier observation lying below the average trend. The reason is that Turkey is an exception in OECD countries, where the gap between female and male employment started to widen in the mid-1990s. In the 1980s, Turkey experienced increasing levels of labor force participation of females similar to those of more developed countries, due to participation of them in agricultural activities in large numbers. However, while most of the OECD countries experienced further increases in the female labor force participation throughout 1990s, especially with increasing role of service sector in income, Turkey experienced the opposite (World Bank and SPO, 2009). FLFPR (FSE) decreased about 30% (15%) in the period 1990-2005 in Turkey (OECD Labour Force Statistics).

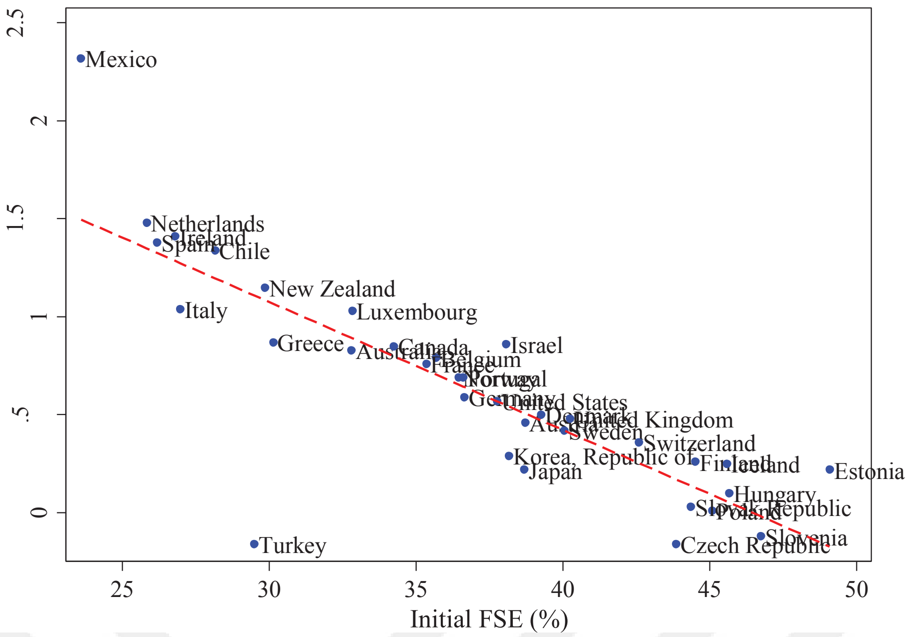

FIGURE 3.2: FSE Convergence of OECD countries, 1971-2010
Data Source: OECD Labour Force Statistics, ILOSTAT- ILO Database of Labour Statistics, Author's calculations.

Supporting the hypothesis also with the descriptive evidence, in this chapter, a dynamic panel convergence process is developed in order to investigate convergence behaviors of the labor market gender equality indicators, FLFPR and FSE. The Difference and the System Generalized Method of Moments (GMM) estimators are employed for the 5-year span unbalanced panel data of 34 OECD countries over the period 1971-2010. Following the income convergence literature, two forms of convergence equations are estimated for each methodology: first, absolute (unconditional) gender equality convergence in the labor market and next, conditional gender equality convergence in the labor market by using various macroeconomic control variables, including GDP per capita, tertiary education, fertility rate, trade openness and foreign direct investment (FDI) inward stock. All GMM estimations reveal strong evidence of absolute and conditional convergence of the labor market gender equality indicators. Furthermore, the estimations show that the (implicit) speed of convergence conditional on each of control variables is higher than that of absolute convergence. In other words, all control variables, which are found to be statistically significant and consistent with the literature, contribute to gender equality convergence in the labor market. All GMM estimations are confirmed to be consistent and robust in terms of the validity of instrument set, and of the expected signs and the significance levels of coefficients of the determinants of the labor market gender equality convergence. This chapter provides clear evidence that convergence in the labor market gender equality as a core social and economic development objective in its own right is a common trend across OECD countries.

The rest of the chapter is as follows: Section 3.2 discusses the methodology and data, and summarizes the empirical findings. Section 3.3 presents the concluding remarks with some policy implications.

3.2 Empirical Analysis

3.2.1 Dynamic Panel Data Model

For the estimation procedure of absolute and conditional gender equality convergence in OECD labor market, dynamic panel data methodology is utilized in the tradition of Islam (1995), Caselli et al. (1996), Bond et al. (2001) and Hoeffler (2002). In particular, Difference GMM estimator proposed by Arellano and Bond (1991), and System GMM estimator proposed by Arellano and Bover (1995) and Blundell and Bond (1998) are adopted to overcome modeling issues such as fixed effects, potential endogeneity of regressors and dynamic panel bias. Although it is a widespread practice to employ Ordinary Least Squares (OLS) levels and Within Groups in the convergence literature, these estimators result in biased and inconsistent estimates in dynamic panel data regressions. For OLS levels estimator, (i) the omission of unobserved time invariant country effects causes estimates to be biased and inconsistent; (ii) the coefficient of the lagged dependent variable tends to be biased upwards, since it is positively correlated with the permanent effects in typical growth regressions (Hsiao, 2014). For Within Groups estimator, which takes into account unobserved country-specific effects, (i) estimates are biased and inconsistent with a fixed time period; (ii) in contrast with OLS levels estimate, the estimate of the coefficient of the lagged dependent variable tends to be biased downwards (Nickell, 1981). The Difference and the System GMM estimators tackle the drawbacks of the two methods (Roodman, 2009a). Furthermore, the GMM estimators yield consistent and efficient parameter estimates in a regression with heteroscedasticity and autocorrelation within individuals, and with independent variables that are not strictly exogenous, meaning they are correlated with past and current realizations of the error term (Roodman, 2009a). In order to overcome the endogeneity problem, the Difference GMM eliminates fixed effects by transforming the data, while the System GMM instruments the lagged dependent variable and/or any other endogenous variables with variables thought uncorrelated with the fixed effects (Nickell, 1981; Roodman, 2009a). The Difference and the System GMM are highly recommended for the estimation of growth models, especially with small time dimension, and relatively large cross-sectional dimension (Blundell and Bond, 1998, 2000; Blundell et al., 2001; Bond et al., 2001).

The Difference GMM estimation starts by transforming all regressors, usually by differencing, and uses the Generalized Method of Moments (Hansen, 1982; Roodman, 2009a). However, the Difference GMM can show poor performance when the series are persistent over time, and the number of time series observations is small (Blundell and Bond, 2000; Hoeffler, 2002). The coefficient of the lagged dependent variable estimated by the Difference GMM tends to be biased downwards towards the Within Groups estimation, since the available instruments are only weakly correlated with the endogenous variables. Blundell and Bond (1998) show that the lagged levels of the explanatory variables are weak instruments for the first-differenced equation, hence compose a system with two sets of equations by using not only lagged levels of the series as instruments in the first-differenced equation, but also lagged differences of the series as instruments in the levels equation, which is called the System GMM. Furthermore, the System GMM is more efficient than the Difference GMM with an additional assumption that the first differences of instruments are uncorrelated with the fixed effects, which in turn allows the inclusion of more instruments (Roodman, 2009a).

In line with the per capita income convergence equation, the following dynamic panel equation is derived in order to estimate gender equality convergence in the labor market:[17]

[17] Please refer to Equation (12) in Islam (1995) and Equation (2) in Caselli et al. (1996) for further details of the dynamic panel data equation on per capita income convergence.

$$GE_{it} = \theta \cdot GE_{i,t-1} + \gamma \cdot Z_{it} + \mu_i + \phi_t + \varepsilon_{it} \tag{3.1}$$

The LHS of Equation (3.1), GE_{it}, represents the labor market gender equality indicator, either FLFPR or FSE, over a 5-year time period. On the RHS, the determinants of gender equality convergence in the labor market are specified. θ is the coefficient of the labor market gender equality indicator of the previous 5-year time period, $GE_{i,t-1}$. This coefficient is expected to be between 0 and 1. In accordance with the convergence idea, $\theta^* = \theta - 1$ is then to be between -1 and 0, which implies that the growth of FLFPR or FSE is faster in countries/periods with a lower initial level of FLFPR or FSE.[18] γ denote the corresponding coefficients of the vector of control variables in a 5-year time span, Z_{it}, which appear only in conditional convergence regressions. In addition, μ_i and ϕ_t measure country-specific (fixed) effects and time-specific intercepts, respectively. ε_{it} denotes the transitory error term that varies across countries i, time periods t. Finally, time dummies are used in regressions, since this will empower the assumption of no correlation across individuals in the idiosyncratic disturbances (Roodman, 2009a).

The vector of control variables, Z_{it}, to estimate conditional convergence include GDP per capita, tertiary education, fertility rate, trade openness and FDI inward stock, all of which affect gender equality in the labor market. As discussed in Section 3.1, several studies argue the positive impact of income (economic growth/development) on the labor market gender equality (Forsythe et al., 2000; Richards and Gelleny, 2007; Duflo, 2012). The literature also investigates the demographical and structural effects, such as educational level and fertility rate on women's labor market activities (Pettit and Hook, 2005). Several studies assert that education has a positive impact on female labor force participation (Psacharopoulos and Tzannatos, 1991; Tansel, 1994, 1996, 2002). The evidence suggests that the supply of women workers has increased due to rising educational levels, which affects both the initial decision to participate and the working hours in the labor market (Goldin, 1990). On the other hand, the literature points out an inverse relationship between fertility rate and female employment (Lehrer and Nerlove,

[18] Please note that θ^* would be the coefficient of $GE_{i,t-1}$ if the dependent variable were in growth form, that is, $GE_{it} - GE_{i,t-1}$.

1986; Goldin, 1990; Ahn and Mira, 2002; Bloom et al., 2009; Cavalcanti and Tavares, 2015). Moreover, child rearing responsibilities is found to have a negative impact on women's activities and working hours in the labor market (Maglad, 1998). In regard to the relation between globalization and the labor market gender equality, the literature argues that openness to trade increases job opportunities for women in export-oriented industries, and provide better economic rights for them (Çağatay and Özler, 1995; Standing, 1999; Neumayer and De Soysa, 2007; Bussmann, 2009). Another argument is that FDI inflows enable female employment in export-oriented industries and manufacturing sector (Joekes and Weston, 1994; Braunstein, 2002).

The first step of the estimation procedure is to eliminate the fixed effects (μ_i) through a first difference transformation:

$$\Delta GE_{it} = \theta \cdot \Delta GE_{i,t-1} + \gamma \cdot \Delta Z_{it} + \Delta\phi_t + \Delta\varepsilon_{it}\ for\ i = 1, \dots, N\ and\ t = 3, \dots, T \quad (3.2)$$

Assuming that the transitory errors are independent across individuals and serially uncorrelated:

$$E(\varepsilon_{it}\varepsilon_{is}) = 0\ for\ t \neq s \quad (3.3)$$

and that the initial conditions provide:

$$E(GE_{i1}\varepsilon_{it}) = 0\ for\ t \geq 2 \quad (3.4)$$

These assumptions indicate that $GE_{i,t-1}$ is predetermined with respect to ε_{it}.

The moment restrictions by Arellano and Bond (1991) for the Difference GMM are given in Equations (3.5) and (3.6):

$$E\left(GE_{i,t-s}\Delta\varepsilon_{it}\right) = 0\ for\ t = 3, \dots, T\ and\ s \geq 2 \quad (3.5)$$

$$E\left(Z_{i,t-s}\Delta\varepsilon_{it}\right) = 0\ for\ t = 3, \dots, T\ and\ s \geq 2 \quad (3.6)$$

The additional moment restrictions by Blundell and Bond (1998) for the System GMM are given in Equations (3.7) and (3.8):

$$E\left[\Delta GE_{i,t-1}(\mu_i + \varepsilon_{it})\right] = 0\ for\ t = 3, \dots, T \tag{3.7}$$
$$E\left[\Delta Z_{i,t-1}(\mu_i + \varepsilon_{it})\right] = 0\ for\ t = 3, \dots, T \tag{3.8}$$

In order to ensure the consistency of the Difference and the System GMM estimations, four key diagnostics should be provided. The first one is that there should be no serial correlation in the error term. Arellano-Bond (1991) test investigates the first and the second order serial correlations in the first-differenced residuals. The second-order correlation in first differences is taken in consideration for the analysis of the first-order serial correlation in levels, since this will detect the correlation between $\varepsilon_{i,t-1}$ in $\Delta\varepsilon_{i,t-1}$ and $\varepsilon_{i,t-2}$ in $\Delta\varepsilon_{i,t-2}$ (Roodman, 2009a). Arellano-Bond test for AR(2) in first differences reports the p-values for the null hypothesis of no second-order serial correlation in the first-differenced residuals. The second condition is that there should be no correlation between the instruments and the error term for the instrument validity. Hansen (1982) test of over-identifying restrictions presents the p-values for the null hypothesis of instrument validity. The third condition is that the additional moment restrictions given in Equations (3.7) and (3.8) should be valid for the consistency of System GMM estimations. The System GMM is more efficient than the Difference GMM provided the additional restrictions are valid (Hoeffler, 2002). The Difference-in-Hansen test reports the p-values for the null hypothesis of the validity of additional moment conditions. The final condition is to satisfy the rule of thumb, meaning the number of instruments is smaller than or equal to the number of groups in a regression to eliminate finite sample bias induced by overfitting (Roodman, 2009b).

3.2.2 Data

The empirical evidence is based on unbalanced panel data of 34 OECD countries and the period 1971-2010.[19] The data of FLFPR in percentage, female employment and total employment in thousands for the age range of 15-64 are obtained from OECD

[19] OECD countries are Australia, Austria, Belgium, Canada, Chile, Czech Republic, Denmark, Estonia, Finland, France, Germany, Greece, Hungary, Iceland, Ireland, Israel, Italy, Japan, Republic of Korea, Luxembourg, Mexico, the Netherlands, New Zealand, Norway, Poland, Portugal, Slovak Republic, Slovenia, Spain, Sweden, Switzerland, Turkey, United Kingdom and the United States of America.

Labour Force Statistics and ILOSTAT- International Labour Organization (ILO) Database of Labour Statistics. The source of data of GDP per capita at constant 2005 US$ in millions is The World Bank DataBank. The data of the highest level attained for tertiary education as percentage of population aged 15 and over are retrieved from Barro and Lee Educational Attainment Dataset (2013). The data for total fertility rate, expressed in births per woman, are retrieved from The World Bank DataBank. The source of data of trade openness at 2005 constant prices in percentage is Penn World Table 7.1 of Heston et al. (2012). The data of the share of FDI inward stock in GDP are extracted from United Nations Conference on Trade and Development (UNCTAD) Statistics. Following Islam (1995) and Caselli et al. (1996), 5-year time intervals are preferred in order to reduce serial correlation problem and to eliminate the cyclical component. Accordingly, 8 data (time) points for each of 34 countries are obtained, e.g., 1975, 1980, ..., 2010. In the analysis, all series are in natural logarithms. Table 3.1 presents the average, the dispersion statistics, the minimum, and the maximum of the series in panel data set.

TABLE 3.1: Descriptive Statistics of 5-year span data, 1971-2010

Variables	Observations	Mean	Std. Dev.	Min.	Max.
FLFPR (%)	207	58.14	12.78	27.3	83.00
FSE (%)	224	40.79	5.82	23.59	50.55
GDP per capita (millions)	244	25117	14542	2484	83461
Tertiary education (% of population aged 15 and over)	272	9.28	5.51	0.72	26.26
Fertility rate (births per woman)	272	1.95	0.69	1.16	6.33
Trade openness (%)	260	61.38	44.94	10.29	313.19
FDI inward Stock (% of GDP)	212	26.99	34.19	1.00	249.71

Note: Std. Dev., Min. and Max. denote standard deviation, minimum and maximum, respectively.

3.2.3 Findings

The panel regression results from estimating Equation (3.1) by one-step Difference and System GMM estimators with 5-year span data of 34 OECD countries over the period 1971-2010 are reported in Tables 3.2-3.5.[20] For Difference GMM, the one-step and two-step estimators are asymptotically equivalent for the special case of spherical disturbances (Arellano and Bond, 1991). In such a case, the two-step

[20] The command “xtabond2” is used in Stata (v.13) for System GMM estimations, and the instrument matrix is collapsed with the command “collapse” available in Stata, as mentioned in Roodman (2009b).

estimator is more efficient (Hoeffler, 2002). For System GMM, the two-step estimator is more efficient than the one-step estimator. Though, Monte Carlo studies indicate that the efficiency gain is small, and that the two-step estimator converges only slowly to its asymptotic distribution. The asymptotic standard errors relating to the two-step GMM estimators tend to be seriously biased downwards in finite samples (Blundell and Bond, 1998; Hoeffler, 2002). Therefore, following Hoeffler (2002), one-step GMM estimations are reported. In Tables 3.2-3.5, column (1) presents the absolute convergence estimations, in which the labor market gender equality indicator of the previous 5-year time period (lagged dependent variable) is the only determinant of the labor market gender equality convergence; columns (2)-(6) show the conditional convergence estimations, which include GDP per capita, tertiary education, fertility rate, trade openness and FDI inward stock, respectively as additional determinants of the labor market gender equality convergence.

In Tables 3.2-3.5, the lagged dependent variable is assumed to be predetermined, and the control variables are regarded as endogenous regressors for all regressions.[21] All regressions include time dummies for 8 data points; however, their coefficients are not presented in order to preserve space. The first row presents $\hat{\theta}$, the estimated coefficient of the lagged dependent variable. This coefficient is expected to be between 0 and 1, implying $\widehat{\theta^*} = \hat{\theta} - 1$ to be between -1 and 0, which is an evidence of gender equality convergence in the labor market.[22] However, the expansion of θ is unknown in Equation (3.1) due to the lack of theoretical background for convergence in the labor market gender equality indicators. Hence, the implied convergence rate cannot be calculated.[23] The only interpretation that can be made is the (implicit) speed of convergence: the lower $\hat{\theta}$, hence the higher $\widehat{\theta^*}$ in absolute value implies the higher speed of convergence. Moreover, $\widehat{\theta^*}$ in absolute value is expected to increase with the inclusion of a control variable in the model. In other words, the speed of conditional convergence is expected to be higher than that of absolute convergence.

[21] The endogeneity issue between the labor market gender equality and the control variables is treated correctly by the Difference and the System GMM estimators.
[22] Please note that $\hat{\theta} = 1$ implies that differences in the labor market gender equality across OECD countries persist over time; and $\hat{\theta} > 1$ yields evidence of divergence of OECD countries in gender equality in the labor market.
[23] Please refer to Equation (12) in Islam (1995) and Equation (6-7) in Caselli et al. (1996) for the theoretical expansion of the coefficient on lagged GDP per capita, and for the calculation of convergence rate implied by this coefficient.

Tables 3.2 and 3.3, respectively, present the Difference and the System GMM estimations of the labor market gender equality convergence by using FLFPR as the dependent variable. The coefficient of lagged FLFPR is between 0 and 1, and statistically significant at 1% level in all Difference and System GMM estimations, which provides a concrete evidence of both absolute and conditional FLFPR convergence of OECD countries in the sample period. Confirming a priori expectations, both the Difference and the System GMM regressions of absolute convergence yield the highest $\hat{\theta}$, 0.805 and 0.840, respectively, hence the lowest speed of convergence (columns (1) of Tables 3.2 and 3.3). The inclusion of control variables one by one in the model results in lower values for $\hat{\theta}$, higher absolute values for $\widehat{\theta^*}$. Notably, the speed of conditional convergence is higher than that of unconditional convergence, as expected. Moreover, both the Difference and the System GMM estimations reveal that all control variables are statistically significant at either 1% or 5% level, and that their signs are consistent with a priori expectations (columns (2)-(6) of Tables 3.2 and 3.3). Hence, it is natural to assert that each of the control variables, namely GDP per capita, tertiary education, fertility rate, trade openness and FDI inward stock, has contributed to FLFPR convergence of OECD countries in the sample period.

TABLE 3.2: Difference GMM Estimations of FLFPR Convergence from a panel of 5-year span data, 1971-2010

Variables and Statistics	**Dependent Variable: ln[FLFPR]**					
	(1)	(2)	(3)	(4)	(5)	(6)
ln[FLFPR(-1)]	0.805^{***} (0.05)	0.626^{***} (0.07)	0.579^{***} (0.06)	0.756^{***} (0.05)	0.669^{***} (0.06)	0.693^{***} (0.05)
ln[GDP per capita]	-	0.105^{***} (0.03)	-	-	-	-
ln[tertiary education]	-	-	0.121^{***} (0.04)	-	-	-
ln[fertility rate]	-	-	-	-0.088^{**} (0.04)	-	-
ln[trade openness]	-	-	-	-	0.063^{**} (0.03)	-
ln[FDI inward stock]	-	-	-	-	-	0.023^{**} (0.01)
Instruments	11	19	4	10	8	6
Groups	34	34	34	34	34	34
Hansen test	0.16	0.21	0.50	0.12	0.12	0.15
AR(2)	0.33	0.16	0.82	0.62	0.26	0.74

Notes: "(-1)" denotes one lag of the corresponding variable. Heteroscedasticity-consistent standard errors are in parentheses. Windmeijer (2005) finite sample correction for standard errors is employed. The superscripts *** and ** denote the statistical significance at 1% and 5% levels, respectively.

TABLE 3.3: System GMM Estimations of FLFPR Convergence from a panel of 5-year span data, 1971-2010

Variables and Statistics	Dependent variable: ln[FLFPR]					
	(1)	(2)	(3)	(4)	(5)	(6)
ln[FLFPR(-1)]	0.840^{***}	0.640^{***}	0.597^{***}	0.747^{***}	0.675^{***}	0.604^{***}
	(0.06)	(0.08)	(0.06)	(0.05)	(0.06)	(0.14)
ln[GDP per capita]	-	0.092^{**}	-	-	-	-
		(0.04)				
ln[tertiary education]	-	-	0.108^{***}	-	-	-
			(0.04)			
ln[fertility rate]	-	-	-	-0.095^{**}	-	-
				(0.04)		
ln[trade openness]	-	-	-	-	0.064^{**}	-
					(0.03)	
ln[FDI inward stock]	-	-	-	-	-	0.020^{**}
						(0.01)
Constant	0.692^{***}	0.578^{**}	1.423^{***}	1.016^{***}	1.091^{***}	1.584^{***}
	(0.24)	(0.29)	(0.20)	(0.21)	(0.21)	(0.54)
Instruments	13	21	14	16	14	12
Groups	34	34	34	34	34	34
Hansen test	0.14	0.35	0.28	0.15	0.20	0.61
Difference-in-Hansen test	0.11	0.58	0.18	0.16	0.17	0.30
AR(2)	0.30	0.17	0.77	0.63	0.24	0.74

Notes: "(-1)" denotes one lag of the corresponding variable. Heteroscedasticity-consistent standard errors are in parentheses. Windmeijer (2005) finite sample correction for standard errors is employed. The superscripts *** and ** denote the statistical significance at 1% and 5% levels, respectively.

Tables 3.4 and 3.5, respectively show the Difference and the System GMM estimations of the labor market gender equality convergence by using FSE as the dependent variable. In all Difference and System GMM estimations, the coefficient of lagged FSE is statistically significant at 1% level, and takes the value between 0 and 1, which proves FSE convergence across OECD countries in both absolute and conditional senses in the sample period. In analogy to the estimations in Tables 3.2 and 3.3, $\hat{\theta}$ is the highest in absolute convergence regressions by the Difference and the System GMM, 0.816 and 0.832, respectively, which imply the lowest speed of convergence (columns (1) of Tables 3.4 and 3.5). The inclusion of control variables one by one in the model lowers $\hat{\theta}$, increases $\widehat{\theta^*}$ in absolute value, hence enhances the implicit convergence rate. Furthermore, both the Difference GMM and System GMM estimations indicate that all control variables are statistically significant at either 1% or 5% level, and that their signs are consistent with the literature (columns

(2)-(6) of Tables 3.4 and 3.5). Therefore, each of the control variables has an accelerating impact on FSE convergence of OECD countries in the sample period.

TABLE 3.4: Difference GMM Estimations of FSE Convergence from a panel of 5-year span data, 1971-2010

Variables and Statistics	Dependent Variable: ln[FSE]					
	(1)	(2)	(3)	(4)	(5)	(6)
ln[FSE(-1)]	0.816*** (0.06)	0.580*** (0.11)	0.600*** (0.08)	0.744*** (0.06)	0.545*** (0.12)	0.549*** (0.11)
ln[GDP per capita]	-	0.108*** (0.04)	-	-	-	-
ln[tertiary education]	-	-	0.072*** (0.02)	-	-	-
ln[fertility rate]	-	-	-	-0.136*** (0.05)	-	-
ln[trade openness]	-	-	-	-	0.053** (0.03)	-
ln[FDI inward stock]	-	-	-	-	-	0.029*** (0.01)
Instruments	4	19	19	10	11	7
Groups	34	34	34	34	34	34
Hansen test	0.54	0.10	0.11	0.18	0.17	0.15
AR(2)	0.55	0.15	0.44	0.85	0.19	0.19

Notes: "(-1)" denotes one lag of the corresponding variable. Heteroscedasticity-consistent standard errors are in parentheses. Windmeijer (2005) finite sample correction for standard errors is employed. The superscripts *** and ** denote the statistical significance at 1% and 5% levels, respectively.

For the consistency check of the Difference and the System GMM estimations in Tables 3.2-3.5, four key diagnostics should be taken into account. Firstly, the p-values of AR(2) statistics show no evidence of the second-order serial correlation in the first-differenced residuals in any of the estimations. Secondly, the p-values of Hansen test of over-identifying restrictions indicate the validity of instruments in all estimations. Thirdly, the p-values of Difference-in-Hansen test imply the validity of additional moment restrictions of all System GMM estimations. In this regard, the System GMM estimations are found to be more efficient than the Difference GMM estimations. Finally, the rule of thumb is satisfied, since the number of groups is larger than the number of instruments in all estimations. Hence, all GMM estimations are consistent and robust, since (i) the instrument set used by the estimators are valid, (ii) the signs and the significance levels of the coefficients of the lagged dependent variable and the control variables are consistent with a priori expectations, and (iii) the Difference GMM estimations are not biased downwards, and are close to the System GMM estimations, meaning the instrumental variables used by the estimators are strong.

TABLE 3.5: System GMM Estimations of FSE Convergence from a panel of 5-year span data, 1971-2010

Variables and Statistics	Dependent Variable: ln [FSE]					
	(1)	(2)	(3)	(4)	(5)	(6)
ln[FSE(-1)]	0.832*** (0.04)	0.661*** (0.07)	0.654*** (0.07)	0.760*** (0.05)	0.550*** (0.12)	0.548*** (0.13)
ln[GDP per capita]	-	0.046** (0.02)	-	-	-	-
ln[tertiary education]	-	-	0.058*** (0.02)	-	-	-
ln[fertility rate]	-	-	-	-0.079** (0.04)	-	-
ln[trade openness]	-	-	-	-	0.052** (0.03)	-
ln[FDI inward stock]	-	-	-	-	-	0.025** (0.01)
Constant	0.652*** (0.16)	0.819*** (0.17)	1.175*** (0.24)	0.959*** (0.19)	1.480*** (0.37)	1.630*** (0.46)
Instruments	8	22	20	17	13	21
Groups	34	34	34	34	34	34
Hansen test	0.11	0.19	0.17	0.16	0.19	0.16
Difference-in-Hansen test	0.70	0.19	0.27	0.22	0.89	0.15
AR(2)	0.61	0.28	0.57	0.61	0.30	0.17

Notes: "(-1)" denotes one lag of the corresponding variable. Heteroscedasticity-consistent standard errors are in parentheses. Windmeijer (2005) finite sample correction for standard errors is employed. The superscripts *** and ** denote the statistical significance at 1% and 5% levels, respectively.

3.3 Concluding Remarks and Policy Implications

This chapter empirically examines whether 34 OECD countries converge in the labor market gender equality in the period 1971-2010. In order to increase the reliability of the analysis, both FLFPR and FSE are used as indicators of the labor market gender equality. Inspired by the income convergence equation, a dynamic panel convergence equation is formulated to analyze the convergence behaviors of these indicators. Particularly, the Difference and the System GMM estimators are utilized, since they provide more consistent and more efficient parameter estimates as compared with other panel data estimators. The consistent and robust GMM estimates from a panel of 5-year span data have evidential value for both the absolute and the conditional gender equality convergence in the labor market across OECD area. In addition, the inclusion of each control variable, namely GDP per capita, tertiary education, fertility rate, trade openness and FDI inward stock in the model yield an even higher speed of convergence. Considering that the control variables are statistically significant and consistent with a priori expectations in the estimations, it is inevitable to affirm that each of them contributes to the labor market gender equality convergence. Based on the empirical evidence, this chapter asserts that OECD countries achieve gender equality convergence in the labor market substantially, which in and of itself is a catalyst of social and economic development.

The results on control variables also lead to a number of policy implications, which have the potential, directly or indirectly, to contribute to the progress of the grand gender convergence in the labor market across OECD countries. First, policy-makers should stimulate (tertiary) education, especially for women, across OECD. Second, it is important to increase the level of trade openness and to promote FDI. The free movement of goods, services and capital will contribute to the grand gender convergence. Third, policy makers should follow family-friendly policies and programs in order to vitiate the negative impact of fertility on gender equality in the labor market. Finally, policy makers should be reminded that female workers may be more vulnerable to downfalls in GDP, and that specific short term policies may be needed to sustain the grand gender convergence in the labor market.

CHAPTER 4

HETEROGENEOUS BEHAVIOR OF 'GENDER KUZNETS CURVE' ACROSS OECD COUNTRIES

4.1 Introduction

There is an extensive body of research that discusses the impact of economic development on gender equality in the labor market in the literature. A considerable number of these studies argue that the response of the labor market gender equality to economic development is U-shaped: while gender equality decreases in the initial stages of economic development, it begins to increase when the country develops beyond a certain threshold, which is the direct translation of "Gender Kuznets Curve" (GKC) hypothesis, cf., Eastin and Prakash (2013).[24] Since the seminal study of Boserup (1970), several authors have tested and verified GKC hypothesis using time series, cross-sectional, and panel data models. For instance, Kottis (1990) finds that the female labor force participation follows a U-shaped trend in the economic development process in Greece during the 1970s and the early 1980s. Psacharopoulos and Tzannatos (1989) make a similar observation for 136 countries during the early 1980s. The findings confirm the hypothesis in such a way that the female labor force participation initially decreases in the period of transition from an agrarian subsistence economy, but begins to increase after a particular level of economic development is achieved. Using data of 70 countries at two different time points, 1965 and 1970, Pampel and Tanaka (1986) find that while development excludes females from the labor force at the initial levels, it helps women take up

[24] Although the U-shaped relationship between gender equality in the labor market (indicated by women's status in the labor market, e.g., the female labor force participation rate) and economic development has been previously discussed in several studies, Eastin and Prakash (2013) is the first to introduce the term 'Gender Kuznets Curve' for a further approximation of the relationship, as mentioned in detail in Footnote 25.

expanded labor market opportunities in the later stages. Goldin (1995) observes the hypothesis also with cross-sectional data of more than 100 countries, and explains the U-shaped relationship between gender equality and economic development as follows: in low-income agrarian economies, women participate in the labor force in large numbers as unpaid family workers. In conjunction with rising levels of income in the early stages of industrialization, the manufacturing sector gains importance in economies, which in turn hamper female labor force due to social norms and employer preferences in patriarchal institutions. In the later stages of economic development, service sector comes into prominence, and increasing educational levels of women facilitate employment in managerial and administrative positions for them. This process is explained by substitution and income effects in the labor supply. The income effect is the change in women's labor supply in return for the change in household income, and the substitution effect is the change in women's labor supply in return for the change in their wages, holding income constant. The declining part of the U-shaped pattern demonstrates that income effect dominates substitution effect, whereas the rising part demonstrates the opposite. In a similar manner, Tansel (2002) investigates the hypothesis by pooling data of 67 provinces of Turkey for the years 1980, 1985 and 1990. The cross-province analysis vindicates the U-shaped relationship between the female labor force participation and level of economic development. Using cross-country data pooled for 1985 and 1990, Çağatay and Özler (1995) also confirm that the relationship between long-term development and the share of women in the labor force is U-shaped. Tam (2011), one of the recent studies in this stream, undertakes dynamic panel data estimation of the hypothesis for 130 countries over 31 years, and also finds that the U-shaped pattern between the female labor force participation and economic development seems to hold.

Although the studies supporting the hypothesis predominate in the literature, there are four alternative arguments regarding the impact of economic development on the female labor market activity. Firstly, some studies argue that the U-shaped pattern may be irrelevant. Durand (1975), for example, asserts that despite the initial decrease in the female labor force participation in agriculture with economic development, the U-shaped pattern of the relationship cannot be generalized for developing countries. Similarly, Steel (1981) finds no evidence of the U-shaped trend of the female labor force participation during the modernization process of Ghana's economy in the 1960s. In contrast, the findings reveal an increase in the female labor force participation with the rapid growth of manufacturing employment in the early stages of development. Standing (1978) argues that the determinants of the female labor force participation are too complex to be described by the U-shaped hypothesis. Secondly, some studies argue for the existence of a positive linear relationship between the labor market gender equality and economic development (Gray et al., 2006; Richards and Gelleny, 2007). Economic development opens doors for women in relatively higher paying nonfarm sectors, promotes them to invest in human capital, and hence improves women's status in the labor market (Weiss et al., 1976; Clark, 1991; Clark et al., 1991; Charles, 1992; Forsythe et al., 2000; Duflo, 2012). Thirdly, Tinker (1976) asserts a negative linear response of the labor market gender equality to economic development: the development strategies may empower patriarchal discriminatory institutions, exclude women from productive work, and prompt them to participate in unskilled and low paying jobs, all of which hinder gender equality. Finally, Eastin and Prakash (2013) make a further approximation that the curvilinear relationship between the labor market gender equality and economic development is not necessarily quadratic, it can be cubic instead. The study suggests that the true relationship between gender equality and economic development is an S-shaped: an increase in gender equality in the early stages of development with the improvements in social and political rights, followed by a plateau or even decrease, due to, for example, the empowerment of discriminatory institutions, and finally, a rise caused by the evolution of new norms and institutions favoring gender equalization in employment.[25] According to this study, previous studies suggesting a U-shaped pattern for the relationship have only been able to capture part of the trend.[26]

[25] Eastin and Prakash (2013) call the S-shaped pattern GKC. Following this study, the cubic approximation was indeed tested; however, the cubic form of income did not work at all due to the preferred empirical methodology in the analyses. Therefore, the quadratic relationship between the labor market gender equality and economic development is examined. This relationship is expected to follow a U-shaped pattern, which will be called GKC in this study.

[26] The S-shaped pattern demonstrates that the level of gender equality rises (+) falls (-) and rises (+) during the economic development process. Then, the second and the third phases of S-shaped GKC coincide with the U-shaped pattern.

In all, it appears that the literature is lacking in arguments on the nature of GKC, particularly, on the potential heterogeneous response of the labor market gender equality to economic development across countries. This research is the first in the literature to reveal the country-specific effects of economic development on gender equality, in addition to the common effect for the whole panel of countries. The second contribution of this research is that the labor market gender equality is indicated by the female share in employment, though the literature preferably uses the female labor force participation rate. According to Eastin and Prakash (2013), the female labor force participation rate approximates a key indicator of women's status in the labor market, and reflects gender equality in economic and social dimensions. However, this research asserts that the female labor force participation rate may not reflect economic dimension of gender equality accurately, since it does not discriminate in the composition of female labor force in terms of working versus actively job hunting. However, the female share in employment directly regards women's participation in activities to produce goods or services for pay or profit, and hence captures the economic dimension as well as the social dimension of gender equality in the labor market.

The aim of this chapter is to investigate the long-run relationship between the labor market gender equality (interchangeably the female share in employment) and economic development for 28 selected OECD countries within the period 1990-2012. The narrowing of the gender gap has been a common characteristic of labor markets across OECD countries over the last three decades, which is due to on the one hand, employment gains for women and, on the other, reductions for men (OECD Employment Outlook, 2002). On average across OECD countries, the gender gap in labor force participation narrowed from 23 percentage points to 13 percentage points in the period 1990-2012 (ILO, IMF, OECD and The World Bank Group, 2014). Nevertheless, the countries vary in timing and the degree to which the narrowing has occurred. There are still wide variations in the gender gap among these countries arising from social and economic factors as well as public policies. For instance, the gender gap in Japanese labor market is 25 percentage points, whereas it is just over 10 percentage points on average across the major advanced economies, and only six percentage points in Sweden (Elborgh-Woytek et al., 2013). In addition, the female participation in the labor market was mainly completed in the 1960s and 1970s in the Nordic countries, whereas it is a more recent experience in countries such as Ireland, the Netherlands and Portugal (OECD Employment Outlook, 2002). The low female employment rates are most pronounced in Greece, Italy, Mexico and Turkey, where less than 50% of women work. In contrast, female employment rates are the highest at over 70% in Nordic countries (OECD, 2008). Hence, it is natural to assert that that the relationship between the labor market gender equality and economic development is not necessarily homogeneous across OECD countries.

This research contributes to the literature by identifying the heterogeneous behavior of the labor market gender equality in response to economic development across OECD countries, in addition to the common behavior for the whole panel of countries. To this end, the Common Correlated Effects Mean Group (CCEMG) estimator proposed by Pesaran (2006), and Augmented Mean Group (AMG) estimator proposed by Eberhardt and Bond (2009) are employed. These methods take account of cross-sectional dependence in panel data and provide both the common and the country-specific estimations. The panel findings indicate, consistently with the majority of the literature, that the labor market gender equality displays a U-shaped pattern in the course of economic development. Although the panel of 28 selected countries is found to support GKC hypothesis in the period 1990-2012, the country-specific estimations vary and even contradict. Particularly, the results indicate that gender equality follows a U-shaped path in response to economic development, and supports the hypothesis for 13 countries: Belgium, Canada, Denmark, Finland, France, Italy, Japan, Luxembourg, Mexico, New Zealand, Norway, Switzerland and Turkey. The relationship between gender equality and economic development is found to imply a reverse trend and contradict the hypothesis for three countries: Iceland, Israel and the United States of America (USA). Nevertheless, no significant relationship is found between gender equality and economic development for 12 countries: Austria, Australia, Chile, Germany, Greece, Ireland, Republic of Korea, the Netherlands, Portugal, Spain, Sweden and United Kingdom (UK). These results indicate that gender equality is not necessarily homogeneous, hence not a direct outcome of economic development (after a threshold level of income), and that continuous monitoring and the necessary intervention by policy makers is needed to further stimulate women's economic participation in the labor market. In addition, gender equalization process may be subject to decoupling, that is, the relationship between levels of gender equalization on the one hand, and employment and income growth on the other may be broken, given that the relationship is either insignificant or in reverse direction for some of the most economically advanced countries.

The rest of the chapter is as follows: Section 4.2 introduces the data and methodology, and reports the empirical findings. Section 4.3 provides the discussion of results and concluding remarks.

4.2 Empirical Analysis

4.2.1 Data

The empirical analysis is based on balanced panel data including annual observations for 28 selected OECD countries over period 1990-2012.[27] The dependent variable is the labor market gender equality that is indicated by the female share in employment. The data of female employment and total employment for the age range of 15-64 in thousands are compiled from Key Indicators of the Labour Market (KILM), 8th edition- International Labour Organization (ILO) Database of Labour Statistics. The data of GDP per person employed at constant 1990 PPP $ are obtained from The World Bank DataBank. Table 4.1 reports the summary statistics of the series in panel data set.

[27] The 28 countries consist of 34 OECD countries excluding Post-Soviet due to the limitations in the availability of data. The sample consist of Australia, Austria, Belgium, Canada, Chile, Denmark, Finland, France, Germany, Greece, Iceland, Ireland, Israel, Italy, Japan, Republic of Korea, Luxembourg, Mexico, the Netherlands, New Zealand, Norway, Portugal, Spain, Sweden, Switzerland, Turkey, UK and USA. The analysis starts from 1990, since it is the earliest year for which the data set for female employment is available for some countries in the sample.

TABLE 4.1: Descriptive Statistics of annual data, 1990-2012

Statistics/Variables	Female share in employment (%)	GDP per person employed ($)
Mean	42.37	40702.23
Standard Deviation	5.13	9949.46
Minimum	23.59	15576
Maximum	49.52	68374
Observations	644	644

4.2.2 Methodology and Findings

The long-run relationship between the labor market gender equality and economic development in the context of GKC hypothesis is estimated with the panel regression model in Equation (4.1)

$$\ln(GE_{it}) = \eta_i + \phi_t + \theta_1 \ln(y_{it}) + \theta_2 [\ln(y_{it})]^2 + \epsilon_{it} \quad (4.1)$$

where GE_{it} represents the labor market gender equality, interchangeably the female share in employment. Economic development is measured in terms of GDP per person employed, y_{it}. The model also includes the quadratic specification of y_{it} to test for a curvilinear relationship between gender equality and economic development. In keeping with GKC hypothesis, the long-run elasticities of gender equality with respect to income and the square of income, respectively, are expected to be $\theta_1 < 0$ and $\theta_2 > 0$. This means in the initial stages of economic development, gender equality decreases until some threshold level of GDP per person employed is reached, after which gender equality begins to increase. The U-shaped GKC implies a 'turning point' GDP per person employed, at which the level of the labor market gender equality is the minimum, given in Equation (4.2):[28]

$$\tau = \exp[-\theta_1/(2\theta_2)] \quad (4.2)$$

[28] Kuznets (1955) hypothesis that explains the inverted U-shaped relationship between income inequality and economic development suggests that the turning point income is where the level of income inequality is at the maximum of the curve. Drawing on Kuznets' hypothesis, GKC considers the relationship between gender equality and economic development, which is expected to be U-shaped. Hence, the turning point income is where the level of gender equality is at the minimum of the U-shaped GKC.

Finally, η_i and ϕ_t denote intercept parameters, which vary across countries i and years t, respectively. ϵ_{it} is the error term.

There is a growing body of research in the literature drawing a conclusion that panel data sets are likely to show cross-sectional dependence, which may arise from economic integration of countries, common shocks (such as financial, political and social shocks), and sometimes unobserved factors that eventually become the part of error (disturbance) term (Pesaran, 2004). Since traditional estimation methods have become inconsistent or inefficient in the presence of cross-sectional dependence, new techniques have been developed in panel data econometrics for stationarity and cointegration analysis and estimation procedure, which take account of cross-sectional dependence.[29]

As testing for cross-sectional dependence in panel data is necessary to decide on the estimation method, the first step of the empirical analysis is cross-sectional dependence (CD) tests to analyze the contemporaneous correlation across countries in the panel.[30] Panel A and Panel B of Table 4.2 report the results of Breusch and Pagan (1980), Pesaran (2004) and Pesaran, Ullah and Yamagata (2008) CD tests for each series and for the model, respectively. The results imply that for both the model with intercept and the model with intercept and trend, Bias-Adjusted CD test statistics reject the null hypothesis of no cross-sectional dependence for gender equality, income and the square of income series, and for the model.[31]

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[29] Assuming that cross-sectional dependence is due to common unobserved components, but that they are uncorrelated with the included regressors, the Fixed-Effects (FE) and Random-Effects (RE) estimators are consistent, although not efficient, and the estimated standard errors are biased. However, if the common unobserved components are correlated with the included regressors, the FE and RE estimators are inconsistent and biased. Please see, for example, De Hoyos and Sarafidis (2006).

[30] Please note that if the time dimension (T) is larger than the cross-sectional dimension (N) in a panel data set, CD_{LM1} test of Breusch and Pagan (1980) can be used to test for cross-sectional dependence. However, if N is larger than T in a panel, just as in this analysis (N=28, T=23), the CD_{LM1} test statistic does not attain desirable statistical properties as it shows considerable size distortions. Hence, CD_{LM1} test is presented only for the matter of completeness. Please see, for example, Pesaran (2004).

[31] Please note also that Bias-Adjusted CD test of Pesaran et al. (2008) is taken into consideration, since it exhibits a finite sample behavior, compared to CD_{LM2} and CD_{LM} tests of Pesaran (2004); it successfully controls the size while maintaining satisfactory power in a panel with exogenous regressors. Bias-Adjusted CD test is consistent even when CD_{LM2} and CD_{LM} tests are inconsistent. Please see, for example, Pesaran et al. (2008).

TABLE 4.2: Cross-sectional Dependence Tests Results

Panel A: For the series

Variables	Tests	Model with intercept		Model with intercept&trend	
		statistics	p-values	statistics	p-values
$\ln(GE_{it})$	CD_{LM1}[a]	561.061	0.000	603.278	0.000
	CD_{LM2}[b]	6.658	0.000	8.193	0.000
	CD_{LM}[c]	1.458	0.072	1.146	0.126
	Bias- Adjusted CD[d]	43.443	0.000	39.839	0.000
$\ln(y_{it})$	CD_{LM1}[a]	631.147	0.000	658.072	0.000
	CD_{LM2}[b]	9.207	0.000	10.186	0.000
	CD_{LM}[c]	1.155	0.124	0.778	0.218
	Bias-Adjusted CD[d]	40.646	0.000	39.063	0.000
$[\ln(y_{it})]^2$	CD_{LM1}[a]	623.606	0.000	652.936	0.000
	CD_{LM2}[b]	8.933	0.000	9.999	0.000
	CD_{LM}[c]	0.798	0.212	0.438	0.331
	Bias-Adjusted CD[d]	40.223	0.000	38.559	0.000

Panel B: For the model

	Tests	statistics	p-values
The Model	CD_{LM1}[a]	1575.334	0.000
	CD_{LM2}[b]	43.547	0.000
	CD_{LM}[c]	21.052	0.000
	Bias-Adjusted CD[d]	6.664	0.000

Notes: *Panel A* of Table 4.2 reports the CD_{LM1} test developed by Breusch and Pagan (1980), CD_{LM2} and CD_{LM} tests developed by Pesaran (2004), Bias-Adjusted CD test developed by Pesaran et al. (2008) for individual variables, and *Panel B* of Table 4.2 reports these tests for the model. [a]CD_{LM1} tests the null of zero correlations in the context with N fixed and $T\rightarrow\infty$. [b]CD_{LM2} tests the null of zero correlations in the context with N and T large. [c]CD_{LM} tests the null of zero correlations in the context with N large and T small. [d]Bias-Adjusted CD tests the null of zero correlations in the case of panel models with strictly exogenous regressors and normal errors. The null hypothesis of CD tests is the absence of cross-sectional dependence.

Given the presence of cross-sectional dependence in the panel, the first generation unit root tests become invalid. Therefore, in order to analyze the stationarity features of the series, cross-sectionally augmented Im-Pesaran-Shin (CIPS) panel unit root test proposed by Pesaran (2007) is employed. CIPS test statistics is the sample averages of the individual cross-sectionally augmented ADF (CADF) statistics. The results of CIPS test for the panel and CADF test for individuals are reported in Tables 4.3 and 4.4, respectively. The CIPS test results indicate the failure to reject the null hypothesis of the presence of unit root for gender equality, income and the square of income series for both the model with intercept and the model with intercept and trend. In other words, all series are found to be non-stationary processes for the panel.

TABLE 4.3: CIPS Panel Unit Root Test Results

Variables	Model with intercept statistics	Model with intercept&trend statistics
$\ln(GE_{it})$	-1.902	-1.797
$\ln(y_{it})$	-1.659	-2.004
$[\ln(y_{it})]^2$	-1.649	-1.990

Notes: The null hypothesis of the test is the presence of unit root in panel data with cross-sectional dependence in the form of common factor dependence. The critical values from Pesaran (2007, p.280-281, Tables 2.b and 2.c for N=30, T=20) are -2.32 (1%), -2.15 (5%), -2.07 (10%) for model with intercept; -2.83 (1%), -2.67 (5%), -2.58 (10%) for model with intercept and trend.

TABLE 4.4: CADF Panel Unit Root Test Results

Country	$\ln(GE_{it})$				$\ln(y_{it})$				$[\ln(y_{it})]^2$			
	Model with intercept		Model with intercept&trend		Model with intercept		Model with intercept&trend		Model with intercept		Model with intercept&trend	
	CADF statistics	lag	CADF statistics	lag	CADF statistics	lag	CADF statistics	lag	CADF statistics	lag	CADF statistics	lag
Australia	-1.002	2	-0.96	2	-1.826	2	-1.692	2	-1.812	2	-1.682	2
Austria	-5.165^{***}	2	-1.677	4	-1.679	2	-3.346	2	-1.652	2	-3.301	2
Belgium	-2.152	3	-2.306	3	-0.999	2	-4.952^{**}	2	-0.943	2	-4.852^{**}	2
Canada	-2.259	2	-2.166	3	-1.227	5	-4.181^{**}	5	-1.231	5	-4.173^{**}	5
Chile	0.084	5	0.027	5	-3.133^{*}	5	-1.624	4	-3.066^{*}	5	-1.631	4
Denmark	-3.680^{**}	2	-4.201^{**}	4	-2.942	2	-1.592	2	-2.988	2	-1.612	2
Finland	-2.144	2	-2.001	3	-3.244^{*}	2	-3.890^{*}	2	-3.194^{*}	2	-3.873^{*}	2
France	-1.968	3	-1.962	2	-2.016	2	-2.529	2	-2.002	2	-2.494	2
Germany	-1.188	2	-0.969	2	-3.299^{*}	2	-3.059	2	-3.276^{*}	2	-3.042	2
Greece	-1.720	2	-1.994	4	-1.760	3	-3.094	2	-1.739	3	-3.087	2
Iceland	-1.029	3	-1.575	3	-2.011	5	-1.072	5	-1.923	5	-1.022	5
Ireland	-2.812	2	-1.965	2	-0.908	6	-0.283	6	-0.929	6	-0.295	6
Israel	-1.482	3	-1.327	3	-0.501	2	-2.190	2	-0.481	2	-2.178	2
Italy	-2.558	2	-1.853	2	-2.094	2	-2.358	2	-2.110	2	-2.376	2
Japan	-1.491	3	-1.257	3	-2.641	2	-2.459	2	-2.592	2	-2.412	2
Korea	-1.734	2	-1.648	2	0.839	6	0.806	6	0.603	6	0.779	6
Luxembourg	-2.514	2	-2.616	2	0.274	2	-2.009	2	0.3	2	-2.007	2
Mexico	-1.543	2	-2.993	2	-1.537	2	-1.578	2	-1.533	2	-1.568	2
Netherlands	-1.648	2	-1.663	2	-1.311	2	-0.737	2	-1.252	2	-0.66	2
N. Zealand	-0.644	3	-0.158	2	-1.509	2	-1.734	2	-1.484	2	-1.710	2
Norway	-0.223	2	-0.223	2	-1.313	2	-0.51	2	-1.331	2	-0.54	2
Portugal	-3.177^{*}	2	-3.075	2	-1.155	3	-1.186	6	-1.161	3	-1.202	6
Spain	-1.432	2	-1.964	2	-1.606	2	-2.298	2	-1.578	2	-2.300	2
Sweden	-2.287	2	-2.296	2	-2.132	3	-2.740	2	-2.099	3	-2.728	2
Switzerland	0.223	4	0.354	4	-0.669	2	-0.263	2	-0.666	2	-0.243	2
Turkey	-1.633	2	-2.160	2	-1.041	2	-0.914	2	-1.034	2	-0.891	2
UK	-3.061^{*}	2	-2.678	3	-1.213	2	-1.478	2	-1.199	2	-1.443	2
USA	-3.012^{*}	2	-2.994	2	-3.786^{**}	3	-3.147	3	-3.797^{**}	3	-3.180	3

Notes: The superscripts ***, ** and * denote the rejection of the null hypothesis of the presence of unit root at 1%, 5% and 10% levels, respectively. The critical values from Pesaran (2007, p.275-276, Tables 1.b and 1.c for N=30, T=20) are -4.35 (1%), -3.43 (5%), -3.01 (10%) for model with intercept; -4.97 (1%), -4.01 (5%); -3.56 (10%) for model with intercept and trend.

After having confirmed the non-stationarity of the variables for the panel, the subsequent step is to test for cointegration among the dependent variable and the regressors. Given the presence of cross-sectional dependence in the panel, the first generation panel cointegration tests also become invalid. Hence, the second generation panel cointegration tests are employed by allowing for the dependence of cross-sectional units. In particular, LM Bootstrap test of Westerlund and Edgerton (2007), and Durbin-Hausman test of Westerlund (2008) are utilized to ensure the presence of cointegration among gender equality and its potential determinant series, income and the square of income. Panel A and Panel B of Table 4.5 show the results of LM Bootstrap and Durbin-Hausman tests, respectively. There is a strong evidence of cointegration among gender equality, income and the square of income, since LM Bootstrap test results indicate the failure to reject the null hypothesis of the presence of cointegration, and Durbin-Hausman test results reveal the rejection of the null hypothesis of no cointegration.

TABLE 4.5: Panel Cointegration Tests Results

	LM statistics	Bootstrap[a] p-values
Panel A: LM Bootstrap Test		
Model with intercept	0.726	0.991
Model with intercept&trend	4.124	0.568
Panel B: Durbin-Hausman Test		
dh_group	6.872	0.000
dh_panel	2.929	0.002

Notes: *Panel A* reports the results of LM bootstrap panel cointegration test developed by Westerlund and Edgerton (2007). [a]The critical value (95%) is based on the bootstrapped distribution with 5000 bootstrap replications. The null hypothesis of the test is the presence of cointegration among $\ln(GE_{it})$ and its potential determinant series, $\ln(y_{it})$ and $[\ln(y_{it})]^2$. The asymptotic p-values for the test are not presented, since they are computed on the assumption of cross-sectional independence. *Panel B* reports the results of Durbin-Hausman panel cointegration test developed by Westerlund (2008). The null hypothesis of the test is the absence of cointegration among $\ln(GE_{it})$ and its potential determinant series, $\ln(y_{it})$ and $[\ln(y_{it})]^2$. The panel statistic, denoted by dh_panel, is obtained by summing n individual terms before multiplying them together, whereas group mean statistic, denoted by dh_group, is obtained by first multiplying the various terms and then summing. The bootstrap critical values are proposed by Westerlund and Edgerton (2007).

Given the evidence of cointegration among the dependent variable and its potential determinant series, the long-run relationship in the panel regression model, given in Equation (4.1), is further estimated by two methods for panel cointegration estimation. The cross-section augmented cointegrating regression for each country is estimated by Common Correlated Effects (CCE) estimator proposed by Pesaran (2006), and Augmented Mean Group (AMG) estimator proposed by Eberhardt and Bond (2009). The latter allows for cross-sectional dependency, which potentially arises from multiple unobserved common factors. The CCE estimation procedure is advantageous, since it enables augmenting the basic regression with cross-section averages of the dependent variable and the observed regressors as proxies for the unobserved common factors. The CCE estimation procedure is presented in Equation (4.3).

$$\ln(GE_{it}) = \alpha_i + \gamma_i X_{it} + \zeta_1 \overline{\ln(GE_{it})} + \zeta_2 \bar{X}_t + \vartheta_{it} \; for \; i = 1, \dots, N \; and \; t = 1, \dots, T \tag{4.3}$$

where the coefficients ζ_1 and ζ_2 represent the elasticity estimates of $\ln(GE_{it})$ with respect to the cross-section averages of $\ln(GE_{it})$ and the observed regressors, respectively. Accordingly, $\ln(y_{it})$ and $[\ln(y_{it})]^2$ are contained in X_{it}. ϑ_{it} denotes the error term. This procedure allows the individual countries to respond to common time effects differently as reflected by the country-specific coefficients on the cross-sectionally averaged variables. It also provides consistent estimates even when the observed regressors are correlated with the common factors. Using this procedure, the individual coefficients, γ_i, can be estimated in a panel framework. The Common Correlated Effects Mean Group (CCEMG) estimation is a simple average of the individual CCE estimations. The CCEMG estimation procedure is shown in Equations (4.4) and (4.5).

$$\hat{\gamma}_{CCEMG} = \sum_{i=1}^{N} CCE_i / N \tag{4.4}$$

$$SE(\hat{\gamma}_{CCEMG}) = \left[\sum_{i=1}^{N} \sigma\left(\hat{\gamma}_{CCE_i}\right)\right] / \sqrt{N} \tag{4.5}$$

where $\hat{\gamma}_{CCEMG}$ and $SE(\hat{\gamma}_{CCEMG})$ are the estimated CCEMG coefficients and their standard deviations, respectively.

On the other hand, the AMG estimator regards time series data properties as well as the differences in the impact of observables and unobservables across panel groups. This estimator takes account of cross-sectional dependence through the involvement of a ‘common dynamic effect’ in the country regression, which is extracted from the year dummy coefficients (D_t) of a pooled regression in first differences (FD-OLS), and represents the levels-equivalent mean evolvement of unobserved common factors across all countries (Eberhardt and Bond, 2009). Provided that the unobserved common factors compose part of the country-specific cointegrating relation, the augmented country regression model embraces the cointegrating relationship that is allowed to differ across countries. In this regard, it coincides with the assumption of CCEMG estimator (Pedroni, 2007; Eberhardt and Bond, 2009). The first stage stands for a standard FD-OLS regression with T-1 year dummies in first differences, from which the year dummy coefficients, relabeled as $\hat{\mu}_t^{\circ}$, are collected. In the second stage, this variable is included in each of the N standard country regressions. Then, the AMG estimations are derived as averages of the individual country estimations. The first and the second stages of AMG estimation procedure are shown in Equations (4.6) and (4.7), respectively.

AMG – Stage (i) $$\Delta \ln(GE_{it}) = \beta' \Delta X_{i,t} + \sum_{t=2}^{T} c_t \Delta D_t + e_{it} \tag{4.6}$$

$$\Rightarrow \hat{c}_t = \hat{\mu}_t^{\circ}$$

AMG – Stage (ii) $$\ln(GE_{it}) = \varphi_i + \beta_i' X_{i,t} + c_i t + d_i \hat{\mu}_t^{\circ} + e_{it} \tag{4.7}$$

$$\hat{\beta}_{AMG} = N^{-1} \sum_i \hat{\beta}_i$$

where φ_i is constant, and e_{it} denotes the error term of stage (i) and stage (ii). $\hat{\beta}_{AMG}$ stands for cross-sectional group-specific AMG estimations which are averaged across the panel.

Eberhardt and Bond (2009) compare the performance of AMG and CCEMG estimators through Monte Carlo simulations, and find robust results for both approaches. On that account, the panel estimations and the country-specific estimations by both approaches are reported in Tables 4.6 and 4.7, respectively.

TABLE 4.6: Mean Group Type Estimations of GKC, 1990-2012

Dependent Variable: $\ln(GE_{it})$		
Variables	*Panel A: CCEMG Estimator*	*Panel B: AMG Estimator*
$\ln(y_{it})$	-14.518 (5.897)***	-15.701 (8.026)**
$[\ln(y_{it})]^2$	0.693 (0.284)**	0.755 (0.396)*
Turning point income ($)	35408.25	32794.40

Notes: The superscripts ***, **and * denote the statistical significance at 1%, 5% and 10% levels, respectively. Asymptotic standard errors are in parentheses.

The panel regression results from estimating Equation (4.1) by the CCEMG and the AMG estimators for 28 OECD countries over the period 1990-2012 are reported in Panel A and Panel B of Table 4.6, respectively. The panel findings indicate that the elasticity estimate of gender equality with respect to income, $\widehat{\theta_1}$, is negative and statistically significant at 1% and 5% levels, and the elasticity estimate of gender equality with respect to the square of income, $\widehat{\theta_2}$, is positive and statistically significant at 5% and 10% levels for CCEMG and AMG estimators, respectively. Hence, both estimators yield evidence of a common non-linear pathway of the labor market gender equality that OECD countries follow in the course of economic development. The estimations reveal that the panel of OECD countries support GKC hypothesis: the response of gender equality to economic development implies a U-shaped pattern ($\theta_1 < 0, \theta_2 > 0$), which is consistent with a priori expectations. The turning point GDP per person employed at which the level of gender equality is the minimum, and after which it begins to increase, is 35408.25 for CCEMG estimator, and 32794.40 for AMG estimator. These values are in the range of the maximum and the minimum, and close to the mean value of GDP per person employed of the panel data set (see Table 4.1). Therefore, GKC implied by OECD countries in the period 1990-2012 more closely resembles a true parabola.

TABLE 4.7: Country-Specific Estimations of GKC, 1990-2012

Country	**Dependent Variable: $\ln(GE_{it})$**			
	CCE Estimator		AMG Estimator	
	$\ln(y_{it})$	$[\ln(y_{it})]^2$	$\ln(y_{it})$	$[\ln(y_{it})]^2$
Australia	0.765 (6.925)	-0.036 (0.324)	-0.023 (5.346)	0.003 (0.251)
Austria	10.008 (8.776)	-0.480 (0.414)	9.349 (6.418)	-0.440 (0.301)
Belgium	-33.160 (11.027)***	1.543 (0.513)***	-23.194 (8.310)***	1.067 (0.383)***
Canada	-12.410 (8.476)	0.574 (0.397)	-12.287 (4.670)***	0.568 (0.218)***
Chile	-7.388 (8.977)	0.376 (0.440)	15.234 (11.779)	-0.759 (0.583)

TABLE 4.7: Continued

Country	Dependent Variable: $\ln(GE_{it})$			
	CCE Estimator		AMG Estimator	
	$\ln(y_{it})$	$[\ln(y_{it})]^2$	$\ln(y_{it})$	$[\ln(y_{it})]^2$
Denmark	-45.229 (15.342)***	2.112 (0.722)***	-38.297 (7.630)***	1.782 (0.357)***
Finland	-12.584 (5.789)**	0.590 (0.275)**	-11.213 (2.839)***	0.517 (0.133)***
France	-48.227 (21.931)**	2.248 (1.012)**	-15.572 (19.495)	0.722 (0.903)
Germany	15.511 (15.519)	-0.743 (0.731)	-16.813 (13.561)	0.791 (0.640)
Greece	4.169 (7.202)	-0.211 (0.349)	9.674 (6.982)	-0.471 (0.338)
Iceland	19.938 (10.110)**	-0.940 (0.476)**	19.813 (7.913)***	-0.944 (0.374)***
Ireland	-5.470 (4.728)	0.261 (0.220)	2.845 (5.115)	-0.137 (0.240)
Israel	37.717 (21.161)*	-1.780 (0.994)*	68.298 (15.001)***	-3.223 (0.708)***
Italy	-13.047 (17.593)	0.600 (0.822)	-48.240 (21.014)**	2.251 (0.982)**
Japan	-3.157 (6.554)	0.155 (0.308)	-12.488 (7.428)*	0.587 (0.350)*
Korea	0.094 (1.187)	0.004 (0.058)	-1.003 (1.355)	0.064 (0.068)
Luxembourg	-63.502 (35.408)*	2.941 (1.629)*	-83.667 (42.565)**	3.857 (1.954)**
Mexico	-99.590 (68.117)	5.093 (3.473)	-179.249 (57.534)***	9.114 (2.928)***
Netherlands	-5.118 (9.527)	0.241 (0.443)	-4.102 (7.375)	0.199 (0.345)
N. Zealand	-17.470 (10.671)	0.840 (0.512)	-13.332 (7.214)*	0.642 (0.346)*
Norway	-6.784 (4.505)	0.319 (0.209)	-7.934 (2.612)***	0.368 (0.121)***
Portugal	-0.628 (6.196)	0.023 (0.304)	1.237 (6.109)	-0.071 (0.301)
Spain	-36.041 (27.879)	1.704 (1.270)	-21.533 (17.423)	1.019 (0.824)
Sweden	-3.310 (3.408)	0.145 (0.159)	-3.280 (2.269)	0.149 (0.107)
Switzerland	-90.508 (21.032)***	4.234 (0.993)***	-69.198 (18.394)***	3.259 (0.870)***
Turkey	-5.078 (6.565)	0.232 (0.328)	-12.120 (3.876)***	0.575 (0.194)***
UK	4.647 (4.081)	-0.210 (0.190)	3.200 (2.926)	-0.154 (0.137)
USA	9.345 (3.277)***	-0.424 (0.150)***	4.264 (3.751)	-0.202 (0.172)

Notes: The superscripts ***, ** and * denote the statistical significance at 1%, 5% and 10% levels, respectively. Asymptotic standard errors are in parentheses.

The country-specific regression results from estimating Equation (4.1) by the CCE and the AMG estimators for 28 OECD countries over the period 1990-2012 are reported in Table 4.7. Although the panel findings confirm the U-shaped relationship between gender equality and economic development, the country-specific findings reveal that the relationship is not necessarily homogenous across individuals. At country level, the signs of significant coefficients by the CCE estimator indicate that the relationship between gender equality and economic development shows a U-shaped trend for Belgium, Denmark, Finland, France, Luxembourg, and Switzerland, but an inverted U-shaped trend for Iceland, Israel and USA. On the other hand, based on the signs of significant coefficients by the AMG estimator, there are more countries that support GKC hypothesis as compared to the CCE estimator; the relationship between gender equality and economic development implies a U-shaped pattern for Belgium, Canada, Denmark, Finland, Italy, Japan, Luxembourg, Mexico, New Zealand, Norway, Switzerland and Turkey. However, the overturn pattern of the curve for Iceland and Israel is also supported by the AMG estimator, since the results reveal an inverted U-shaped trend of gender equality in economic development process for these two countries. Neither the CCE nor the AMG estimator yields significant relationship between gender equality and economic development for Austria, Australia, Chile, Germany, Greece, Ireland, Republic of Korea, the Netherlands, Portugal, Spain, Sweden and UK.

4.3 Concluding Remarks and Discussions

This chapter empirically investigates whether 28 selected OECD countries confirm GKC hypothesis in the period 1990-2012. GKC hypothesis argues that the labor market gender equality follows a U-shaped path in response to economic development. The empirical evidence for a single country and/or a panel of countries, however, does not always support this argument. In this regard, the main contribution of this research is to capture the heterogeneous behavior of gender equality (interchangeably the female share in employment) in response to economic development across OECD countries. To this end, the CCEMG estimator proposed by Pesaran (2006) and the AMG estimator proposed by Eberhardt and Bond (2009) are utilized. These estimators account for cross-sectional dependence in the panel, and determine the country-specific effects of economic development on gender equality, in addition to the common effect for the whole panel of countries. The findings indicate that although the panel of the 28 countries supports GKC hypothesis in the period 1990-2012, the country-specific responses of gender equality to economic development vary and even contradict. In particular, the results reveal that the relationship between gender equality and economic development implies a U-shaped pattern, and supports the hypothesis for 13 countries (Belgium, Canada, Denmark, Finland, France, Italy, Japan, Luxembourg, Mexico, New Zealand, Norway, Switzerland and Turkey), whereas it shows a reverse trend, and contradicts the hypothesis for three countries (Iceland, Israel and USA). Moreover, the response of gender equality to economic development is not found to be statistically significant for 12 countries (Austria, Australia, Chile, Germany, Greece, Ireland, Republic of Korea, the Netherlands, Portugal, Spain, Sweden and UK). These findings indicate that gender equalization process is not necessarily homogenous, and not a direct outcome of economic development (after a threshold level of income). Therefore, enlightened government policies and labor market interventions are needed to put gender issues on the right track even in the most developed economies. Given that the relationship between gender equality and economic development is either insignificant or in reverse direction for some of the most developed economies, there is room for speculation on the cause of the heterogeneous behavior of gender equality. It is possible that gender equality starts to lag behind economic development, and may even stop after a certain threshold income level is reached. This raises the issue of gender equality decoupling, which in and itself is a subject for future research.

CHAPTER 5

THE NATURE AND INCOME ELASTICITY OF GENDER EQUALITY IN THE LABOR MARKET: THE G7 CASE

5.1 Introduction

Most developed economies have experienced major improvements in women's labor market activities, helping to increase gender equality in terms of the labor force measures, especially in recent decades (OECD Employment Outlook, 2015). This is clearly seen in the example of G7 countries, the most advanced economies in terms of income and social transformation.[32] In the period 1984-2014, the increases in FLFPR (FSE) are 20.6% (13.7%) in Canada, 18.9% (16.7%) in France, 42.9% (21.5%) in Germany, 38.4% (30.2%) in Italy, 21% (9.5%) in Japan, 16.9% (13.3%) in UK, and 6.7% (7.2%) in USA (OECD Labour Force Statistics). Nonetheless, the narrowing of the gender gap in labor force participation rate, employment and/or other measures of the labor market has not been sufficient to indicate the completion of the process of "grand gender convergence in labor market", cf., Goldin (2014). The 2007/2008 global financial crisis is a good illustration of the potential for stagnation, and even reversal of the process. If FLFPR and FSE series are analyzed for the period 2007-2014, the annual average increases slow dramatically in all countries but Japan. Since 2007, for example, the increase in FLFPR (FSE) was only 0.08% (0.89%) in Canada and -2.83% (1.14%) in USA (OECD Labour Force Statistics). UN Women (2014) report similarly observes that 'the gender gap in employment has become higher even than its pre-crisis level in most of the advanced economies in the period 2009-2012'.

[32] G7 countries are Canada, France, Germany, Italy, Japan, United Kingdom (UK) and the United States of America (USA).

The negative effect of the recent global financial crisis on gender equality in the labor market is just one factor highlighting that the gender equality process can be subject to stagnation and even reversal. If devolution is possible for the most developed economies, furthermore, the process is even more exposed to reversals in developing economies. In this regard, it is important to identify two conditions: (i) whether the labor market gender equality series are shock resistant, (ii) if and to what extent these series are affected by income, considering the shocks on them. The first is to ascertain whether the impacts of shocks on gender equality in the labor market are permanent or transitory in nature, by additionally allowing for structural breaks in the process. The second is to determine long-run and short-run income elasticities of the labor market gender equality series, considering structural breaks. These analyses combined provide policy makers valuable information in understanding the extent to which the labor market gender equality series are resistant to shocks, especially during economic downturns. This is the major contribution of this research.

The female labor force participation rate (FLFPR) is often considered as the barometer of women's status and gender equality, since it captures the social and economic improvement in the labor market (Eastin and Prakash, 2013). Since the seminal study of Mincer (1962) and Cain (1966), female labor force participation as an indicator of women's labor market activity has been incorporated into several economic analyses , including those of G7 countries, e.g., Trovato and Vos (1992) for Canada, Riboud (1985) for France, Börsch-Supan and Schnabel (1998) and Franz (1985) for Germany, Colombino and De Stavola (1985) for Italy, Shimada and Higuchi (1985) and Sasaki (2002) for Japan, Viitanen (2005) for UK, Oppenheimer (1970), Lehrer and Nerlove (1986) and Antecol (2000) for USA. Nevertheless, FLFPR may not accurately reflect the situation in regard to gender equality in the labor market, as it ignores the classification in the composition of female labor force in the sense of being employed or unemployed, but seeking work. Therefore, the female share in employment (FSE) is utilized as a second indicator of the labor market gender equality to eliminate a statistical artifact due to measurement variation and to increase the reliability of the analysis. The reason is that FSE directly counts in active engagement of women in economic activity in producing goods and providing services, unlike FLFPR.

The literature on gender is lacking in studies about the persistence in time series. This research is the first to investigate the unit root properties of FLFPR and FSE series, by also regarding endogenously determined structural breaks. It is known from time series econometrics that the impact of a shock on a series is temporary if it is mean-reverting. If not, the mean or variance of the series will change over time, and the shock will have permanent impact. The analysis of univariate time series structure of macroeconomic variables started with the early study of Nelson and Plosser (1982), which applied Dickey- Fuller type test (Dickey, 1976; Fuller, 1976; Dickey and Fuller, 1979) to 14 annual time series for USA, and failed to reject the unit root in all but one of them. Since then, the unit root hypothesis has become a topic of great interest in applied macroeconomic studies, e.g., Stock and Watson (1986), Campbell and Mankiw (1987, 1989), Schwert (1987), Perron and Phillips (1987), Evans (1989), Kwiatkowski et al. (1992), Cheung and Chinn (1996), Murray and Nelson (2000), Sollis, Newbold and Leybourne (2000), Lettau and Ludvigson (2004). However, Perron (1989) indicated that the failure to allow for an existing break leads to a bias, reducing the ability to reject a false unit root null hypothesis. In this regard, Perron (1989) highlighted the importance of structural breaks when testing for unit root processes, and proposed a test that would allow for a known or exogenous structural break in the augmented Dickey-Fuller (ADF) tests. Many subsequent studies, such as Banerjee, Lumsdaine and Stock (1992), Zivot and Andrews (1992), Perron and Vogelsang (1992) and Perron (1997) suggested determining the break point endogenously from data. Lumsdaine and Papell (1997) extended Zivot and Andrews (1992) by allowing two endogenous structural breaks in data series. Lee and Strazicich (2003) showed that these ADF type tests have tendency for spurious rejections in finite samples when a break is present under the null hypothesis, since they either do not allow for a break under the null (Zivot and Andrews, 1992 and Lumsdaine and Papell, 1997 tests), or include the break as an innovational outlier (Perron, 1997 test). Hence, Lee and Strazicich (2003) proposed a two break minimum Lagrange Multiplier (LM) test, in which the alternative hypothesis unambiguously implies that the series is trend stationary. However, Popp (2008) indicated that these spurious rejections are not a common characteristic in ADF type unit root tests, and highlighted that the source of spurious rejections is rather the parameters of the test regression having different interpretations under the null and alternative hypotheses, which is of key importance due to their implications for the selection of the structural break date. Accordingly, Narayan and Popp (NP) (2010) overcame this problem by formulating the data generating process as an unobserved components model, and developed an ADF type unit root test for the case of innovational outliers. The test allows for two structural breaks at unknown times under both the null and the alternative hypotheses.[33]

This chapter also contributes to the literature by determining long-run and short-run income elasticities of FLFPR and FSE, taking account of structural breaks in these series. Two main perspectives are held on the (linear) impact of income on gender equality in the labor market, as discussed in several studies. The first perspective, by which the mainstream literature is dominated, suggests that gender equality increases in correspondence with increasing levels of income (Gray et al., 2006; Richards and Gelleny, 2007). Economic development boosts women's labor market position, since it creates employment opportunities for women in relatively higher paying nonfarm sectors, and encourages women to show overall production skills (Weiss et al., 1976; Clark, 1991; Clark et al., 1991; Charles, 1992; Forsythe et al., 2000; Duflo, 2012). In contrast, the second perspective asserts that economic development may strengthen patriarchal discriminatory institutions that isolate women from productive and high paying labor market activities, which in turn undermine gender equality (Tinker, 1976). The labor market conditions should be reorganized in the course of economic development to overcome the entrenched structure of patriarchal institutions (Kabeer, 1996).

The empirical analysis covers G7 countries and the period 1984-2014, due to the limitations in the availability of FLFPR and FSE data. In the first step of analysis, the stationarity features of FLFPR and FSE series are examined using the recently developed NP (2010) test, which allows for two structural breaks in the process. The

[33] Following these developments, several studies investigate unit root properties of key macroeconomic variables considering one or more structural breaks, e.g., Takeuchi (1991), Raj (1992), Ben-David and Papell (1995), Soejima (1995), Li (2000), Mehl (2000), Aguirre and Ferreira (2001), Ben-David, Lumsdaine and Papell (2003), Smyth and Inder (2004), Narayan and Narayan (2010). As this research is concerned with individual countries by time series analyses, the panel unit root tests with structural breaks are not undertaken. However, it is possible to find in the literature a body of research employing panel unit root tests with structural breaks for the stationarity analysis of macroeconomic variables, e.g., Narayan (2004), Narayan (2008a, 2008b).

results indicate that both FLFPR and FSE are characterized by non-stationary behaviors, implying that the shocks have permanent impacts on gender equality in the labor market in all G7 countries. In the second step, the long-run and short-run income elasticities of FLFPR and FSE are estimated by additionally taking into account identified structural break dates of the series through autoregressive distributed lag (ARDL) bounds testing approach of cointegration. The long-run estimates indicate that income has a positive impact on FLFPR in Canada, Germany, UK and USA, on FSE in also France. Furthermore, the short-run analysis indicates that, after any deviation, there is correction and return to the long-run equilibrium level in each period. In all, the findings reveal that a negative shock on income tends to produce downfall impacts on gains in gender equality in the labor market. Hence, the most effective approach to dealing with the periods of fall could be structural policies and institutional adjustments in the labor market aimed at maintaining the momentum in gender equality improvement in the labor market of G7 countries.

The rest of the chapter is as follows: Section 5.2 discusses data. Section 5.3 introduces empirical methodology, and presents the findings. Finally, a summary and some concluding remarks are provided in Section 5.4.

5.2 Data

The annual time series analysis is based on balanced data of gender equality in the labor market (*GE*) and income (*y*) for G7 countries in the period 1984-2014. In order to ensure the reliability of the analyses, FLFPR and FSE are used alternatively to indicate the labor market gender equality. The data of FLFPR in percentage, female employment and total employment in thousands for the working age are compiled from OECD Labour Force Statistics. The data of GDP per capita at constant 2005 US$ in millions are from The World Bank DataBank. All series are in natural logarithms in the empirical analyses. Table 5.1 reports the average, the dispersion statistics, the minimum and the maximum of the series for the seven countries.

TABLE 5.1: Descriptive Statistics of annual data, 1984-2014

	FLFPR (%)				FSE (%)				GDP per capita (millions)			
Country	Mean	S.D.	Min.	Max.	Mean	S.D.	Min.	Max.	Mean	S.D.	Min.	Max.
Canada	70.07	3.72	61.56	74.66	46	1.75	42.2	48.31	31908.18	4382.59	25218.35	38293.28
France	61.99	3.45	56.68	67.37	44.97	2.12	41.32	48.24	31441.44	3912.53	24348.55	36073.52
Germany	62.97	6.65	51.02	72.91	43.11	2.74	38.52	46.8	32552.79	4474.42	24390.25	39717.7
Italy	46.77	4.63	39.86	55.18	37.07	3.13	32.46	42.27	28721.93	3029.41	22142.4	32830.73
Japan	59.53	3.07	54.47	65.97	41.24	1.01	39.64	43.43	33012.55	3800.1	23650.17	37595.18
UK	68.05	2.55	61.72	72.15	45.29	1.68	41.46	47.08	33961.79	5776.4	23722.19	41567.28
USA	68.41	1.93	62.92	70.73	46.15	0.92	43.83	47.44	38512.7	5788.64	28416.49	46405.26

Note: S.D., Min. and Max. denote standard deviation, minimum and maximum, respectively.

5.3 Methodology and Findings

5.3.1 Unit Root Test

The stationarity features of the series are analyzed through unit root test. Macroeconomic time series may be exposed to structural break(s) due to structural changes in economies, such as financial shocks, regime shifts, and policy changes, all of which are either exogenous or endogenous to the series itself. If structural changes are not integrated in the unit root test specification, even when they are present in data generating process, results may be biased towards weakened non-rejection of the non-stationarity hypothesis (Perron, 1989). Hence, considering the movements in income in the period 1984-2014, which include the 2007/2008 crisis, this study focuses on a unit root test that allows for structural breaks in the process. Although it is acknowledged that Zivot and Andrews (1992) test, which allows for one structural break, is an advance over traditional unit root tests, it is argued that the test may lose power in case of two or more breaks (Lee and Strazicich, 2003). To this end, this study employs NP (2010) unit root test, which allows for two endogenous structural breaks under both the null and alternative hypotheses.[34] An additional feature of the test is that the breaks are modeled as innovational outliers, and hence take effect gradually. NP unit root test utilizes Dickey-Fuller strategy, and comprises two different models; Model A allows for two breaks in the level of a trending data series, whereas Model C allows for two breaks in the level and slope of a trending

[34] The simulations of NP (2010) unit root test are carried by GAUSS 10.0.3.

data series. The data generating process of the test is considered as an unobserved components model, as follows:

$$x_t = d_t + \rho u_{t-1} + \epsilon_t \tag{5.1}$$
$$\epsilon_t = \psi^*(L)e_t = A^*(L)^{-1}B(L)e_t$$

In Equation (5.1), d_t is a deterministic component, and u_t is a stochastic component, which are two components of a time series, x_t. Model A and Model C in Equations (5.2) and (5.3), respectively, differ in the determinations of the deterministic component.

$$d_t^{Model\,A} = \alpha + \beta_t + \psi^*(L)\,(\theta_1 DU'_{1,t} + \theta_2 DU'_{2,t}) \tag{5.2}$$
$$d_t^{Model\,C} = \alpha + \beta_t + \psi^*(L) + (\theta_1 DU'_{1,t} + \theta_2 DU'_{2,t} + \gamma_1 DT'_{1,t} + \gamma_2 DT'_{2,t}) \tag{5.3}$$

where $DU'_{i,t} = 1(t > T'_{B,i})$ and $DT'_{i,t} = 1\left(t > T'_{B,i}\right)(t - T'_{B,i})$, $i = 1, 2$. $T'_{B,i}$ for $i = 1, 2$ are the dates of structural breaks, and θ_i and γ_i for $i = 1, 2$ denote the magnitudes of level and slope breaks, respectively.

The test equations for Model A and Model C are presented in (5.4) and (5.5), respectively.

$$x_t^{Model\,A} = \rho x_{t-1} + \alpha_1 + \beta^* t + \theta_1 D(T'_B)_{1,t} + \theta_2 D(T'_B)_{2,t} + \delta_1 DU'_{1,t-1} + \delta_2 DU'_{2,t-1} + \textstyle\sum_{j=1}^{k} \beta_j\,\Delta x_{t-j} + e_t \tag{5.4}$$

where $\alpha_1 = \psi^*(1)^{-1}[(1-\rho)\alpha + \rho\beta] + \psi^{*\prime}(1)^{-1}(1-\rho)\beta$, $\psi^{*\prime}$ is the mean lag, $\beta^* = \psi^*(1)^{-1}(1-\rho)\beta$, $\delta_i = -(\rho - 1)\theta_i$, and $D(T'_B)_{i,t} = 1(t = T'_{B,i} + 1)$, $i = 1, 2$.

$$x_t^{Model\,C} = \rho x_{t-1} + \alpha^* + \beta^* t + \kappa_1 D(T'_B)_{1,t} + \kappa_2 D(T'_B)_{2,t} + \delta_1^* DU'_{1,t-1} + \delta_2^* DU'_{2,t-1} + \gamma_1^* DT'_{1,t-1} + \gamma_2^* DT'_{2,t-1} + \textstyle\sum_{j=1}^{k} \beta_j\,\Delta x_{t-j} + e_t \tag{5.5}$$

where $\kappa_i = \theta_i + \gamma_i$, $\delta_i^* = \gamma_i - (\rho - 1)\theta_i$, and $\gamma_i^* = -(\rho - 1)\gamma_i$, $i = 1, 2$.

In order to test the unit root null hypothesis of $\rho = 1$ against the alternative hypothesis of $\rho < 1$, the t -statistics of $\hat{\rho}$, denoted $t_{\hat{\rho}}$ is used in Equations (5.4) and (5.5). In the first step, there is an examination of a single break, which is selected according to the maximum absolute t-value of the break dummy coefficient θ_1 for Model A, and κ_1 for Model C. The restrictions $\theta_2 = \delta_2 = 0$ and $\kappa_2 = \delta_2^* = \gamma_2^* = 0$ are imposed for Model A and Model C, respectively. Hence, the estimated first break dates for Model A and Model C are presented in Equations (5.6) and (5.7).

$$\hat{T}_{B,1} = \arg\max_{T_{B,1}} \left| t_{\hat{\theta}_1}(T_{B,1}) \right| \; for\ Model\ A \tag{5.6}$$

$$\hat{T}_{B,1} = \arg\max_{T_{B,1}} \left| t_{\hat{\kappa}_1}(T_{B,1}) \right| \; for\ Model\ C \tag{5.7}$$

In the second step, maximizing the absolute t-value of the break dummy coefficient θ_2 for Model A, and κ_2 for Model C, the second break dates for Model A and Model C are estimated as in Equations (5.8) and (5.9).

$$\hat{T}_{B,2} = \arg\max_{T_{B,2}} \left| t_{\hat{\theta}_2}(\hat{T}_{B,1}, T_{B,2}) \right| \; for\ Model\ A \tag{5.8}$$

$$\hat{T}_{B,2} = \arg\max_{T_{B,2}} \left| t_{\hat{\kappa}_2}(\hat{T}_{B,1}, T_{B,2}) \right| \; for\ Model\ C \tag{5.9}$$

Table 5.2a-5.2g present NP unit root test results for FLFPR, FSE, and income series of G7 countries. The results from both Model A and Model C reveal that the unit root null hypothesis is rejected in first differences of FLFPR and FSE series, and in levels of income series for all seven countries, though at different significance levels. That is, real GDP per capita level is a stationary process. In contrast, FLFPR and FSE series are characterized by non-stationary behaviors. If a series is a unit root process, and is exposed to a negative shock, its reversion to a mean value or to a stable path is unlikely. Hence, the impacts of shocks on FLFPR and FSE are permanent for all G7 countries in period 1984-2014.

TABLE 5.2a: NP Unit Root Test Results for Canada

Panel A: Level	**Model A: level**	**Model C: level&slope**
FLFPR	-2.921(3)[2004, 2006]	-4.207(3)[1990, 2001]
FSE	-2.107(3)[1989, 2009]	-4.125(0)[1990, 2008]
y	-4.807(1)**[1989, 2008]	-5.995(1)***[1996, 2005]
Panel B: FD	**Model A: level**	**Model C: level&slope**
FLFPR	-5.777(0)***[1989, 1995]	-7.647(0)***[1990, 2002]
FSE	-6.693(0)***[1991, 2009]	-6.282(0)***[1991, 2007]

TABLE 5.2b: NP Unit Root Test Results for France

Panel A: Level	**Model A: level**	**Model C: level&slope**
FLFPR	-3.157(0)[1993, 2002]	-3.461(0)[1992, 2003]
FSE	-4.021(0)[1990, 2002]	-3.732(0)[1992, 2002]
y	-4.593(1)**[1997, 2008]	-5.303(2)**[2006, 2009]
Panel B: FD	**Model A: level**	**Model C: level&slope**
FLFPR	-6.951(0)***[1995, 2004]	-7.203(0)***[1995, 2004]
FSE	-5.999(3)***[1989, 2001]	-6.205(0)***[1993, 2001]

TABLE 5.2c: NP Unit Root Test Results for Germany

Panel A: Level	**Model A: level**	**Model C: level&slope**
FLFPR	-2.760(0)[1990, 2005]	-4.705(0)[1990, 2004]
FSE	-2.417(0)[1990, 1998]	-3.431(0)[1990, 2002]
y	-5.419(0)**[1989, 2002]	-6.170(0)***[1989, 2008]
Panel B: FD	**Model A: level**	**Model C: level&slope**
FLFPR	-8.209(0)***[1989, 1991]	-9.183(0)***[1989, 1994]
FSE	-5.873(0)***[1989, 1991]	-8.839(0)***[1989, 1993]

TABLE 5.2d: NP Unit Root Test Results for Italy

Panel A: Level	**Model A: level**	**Model C: level&slope**
FLFPR	-2.859(2)[1992, 2000]	-4.423(0)[1992, 2008]
FSE	-3.034(0)[1992, 2003]	-4.561(0) [1992, 2003]
y	-4.844(0)**[2007, 2011]	-5.444(0)**[1991, 2005]
Panel B: FD	**Model A: level**	**Model C: level&slope**
FLFPR	-5.388(0)***[1991, 1995]	-6.930(0)***[1991, 1998]
FSE	-5.721(0)***[1991, 1994]	-7.275(0)***[1991, 2003]

TABLE 5.2e: NP Unit Root Test Results for Japan

Panel A: Level	**Model A: level**	**Model C: level&slope**
FLFPR	-2.767(1)[1998, 2001]	-3.990(1)[1993, 2001]
FSE	-3.895(1)[1994, 2011]	-4.392(0)[1992, 2003]
y	-6.113(0)***[1997, 2007]	-5.726(0)**[1989, 2008]
Panel B: FD	**Model A: level**	**Model C: level&slope**
FLFPR	-5.192(0)**[1991, 1996]	-5.692(1)**[1990, 2001]
FSE	-7.398(0)***[1990, 2005]	-8.708(0)***[1990, 1994]

TABLE 5.2f: NP Unit Root Test Results for UK

Panel A:Level	**Model A: level**	**Model C: level&slope**
FLFPR	-1.981(2)[2002, 2008]	-3.999(0)[1989, 2009]
FSE	-1.721(1)[1990, 2001]	-4.426(1)[1991, 2008]
y	-4.711(1)**[1997, 2008]	-6.646(1)***[1997, 2006]
Panel B:FD	**Model A: level**	**Model C: level&slope**
FLFPR	-5.692(0)***[1989, 2001]	-6.059(0)***[1989, 2005]
FSE	-8.653(0)***[1989, 1992]	-8.323(0)***[1989, 1996]

TABLE 5.2g: NP Unit Root Test Results for USA

Panel A:Level	**Model A: level**	**Model C: level&slope**
FLFPR	-3.682(0)[1991, 2008]	-4.778(0)[1994, 2000]
FSE	-3.614(0)[2003, 2010]	-4.071(2)[1995, 2003]
y	-4.908 (1)**[1997, 2008]	-5.192(0)**[1990, 2004]
Panel B:FD	**Model A: level**	**Model C: level&slope**
FLFPR	-4.777(3)***[1999, 2003]	-5.186(3)**[1990, 2004]
FSE	-7.019(0)***[2006,2009]	-7.507(3)***[2002, 2007]

Notes: FD stands for first difference. The null hypothesis of the test is the existence of unit root with endogenous structural breaks at two time points. Maximum lags of annual data are set to 3, and optimal lags in parentheses are selected by the significance of t-statistics as to the 10% critical values of 1.645. Trimming [0.10, 0.90]. $T_{B,1}$ and $T_{B,2}$ are reported in brackets. The superscripts *** and ** denote the rejection of the null hypothesis at 1% and 5% levels, respectively. The critical values from NP (2010, p.1429, Table 3 for T=50) are -5.259 (1%), -4.514 (5%), -4.143 (10%) for Model A; -5.949 (1%), -5.181 (5%), -4.789 (10%) for Model C.

5.3.2 ARDL Cointegration Analysis

The long-run relationship between the labor market gender equality and income is investigated with Equation (5.10).

$$GE_t = \varphi + \lambda y_t + \tau D_{1t} + \omega D_{2t} + \eta_t \tag{5.10}$$

where GE_t represents the labor market gender equality indicator, alternatively, FLFPR and FSE, and y_t denotes real GDP per capita. φ is constant, and the coefficient λ is the long-run income elasticity of gender equality in the labor market. D_{1t} and D_{2t} are dummy variables, indicating the structural break dates ($T_{B,1}$, $T_{B,2}$) in FLFPR and FSE series.[35] η_t stands for the disturbance term, and the subscript t is the time period index.

The time series analyses are carried out using the ARDL bounds testing approach of cointegration suggested by Pesaran and Shin (1999) and Pesaran, Shin and Smith (2001). ARDL cointegration approach has three main advantages compared to other cointegration models, such as residual-based Engle and Granger (1987) test, and the maximum likelihood test of Johansen (1988) and Johansen and Juselius (1990). The first is that there is no necessity for the variables in the equation to be of the same

[35] $D_1 = 1$ if $t \geq T_{B,1}$, 0 otherwise; $D_2 = 1$ if $t \geq T_{B,2}$, 0 otherwise. $T_{B,1}$ and $T_{B,2}$ for FLFPR and FSE series in first differences are for the level model (Model A) of NP unit root test.

order of integration. The model is suggested to incorporate I(0) and I(1) variables in the same equation.[36] The second is that it allows that the variables may have different optimal lags. The final advantage is that it yields efficient parameter estimates irrespective of the size of the sample and whether the explanatory variables are endogenous. The model corrects the endogeneity problem of explanatory variables even in small samples (Menyah and Wolde-Rufael, 2010).[37]

ARDL is a dynamic single equation regression model, which includes the lagged values of the dependent variable and the current and the lagged values of the regressors, for the direct estimation of short-run elasticities, and indirect estimation of long-run equilibrium relationship (Wang et al., 2011).[38] ARDL bounds testing approach of cointegration is performed in two phases, once it is ensured that all variables are either I(0) or I(1).[39] The first procedure is to test the existence of long-run relationship (cointegration) among dependent variable and regressors through the bounds F-test for cointegration. If the cointegration among variables is determined, then the subsequent procedure is to estimate long-run and short-run models through ARDL approach and Error-Correction Model (ECM), respectively.

To this end, the bounds F-test for cointegration is employed as a first step to detect the existence of long-run relationship between gender equality in the labor market and real GDP per capita through Equation (5.11).

$$\Delta GE_t = \varphi_1 + \sum_{m=1}^{n} \varsigma_{1m} \Delta GE_{t-m} + \sum_{p=0}^{r} \lambda_{1p} \Delta y_{t-p} + \zeta_1 GE_{t-1} + \zeta_2 y_{t-1} + \tau_1 D_{1t} + \omega_1 D_{2t} + \eta_{1t} \quad (5.11)$$

[36] However, the integration order of any of the series should not be greater than one, e.g., I(2), for the applicability of critical bounds provided by Pesaran et al. (2001) or Narayan (2005).
[37] The bidirectional relationship between income and gender equality in the labor market is the main reason for choosing ARDL cointegration approach. The response of income (economic growth/development) to the labor market gender equality in specific measures is also discussed in various studies in the literature, e.g., Tzannatos (1999), Klasen (1999), Esteve-Volart (2004), Stotsky (2006), Morrison et al. (2007), Bloom et al. (2009), Klasen and Lamanna (2009), Löfström (2009), Ferrant (2015).
[38] See Pesaran and Pesaran (1997) for further discussion of the model and an application by MICROFIT econometric software.
[39] NP unit root test results reveal that FLFPR and FSE series are stationary in first differences, i.e., I(1), and that real GDP per capita series are stationary in levels, i.e., I(0) for both Model A and Model C for all countries (please see Table 5.2a-5.2g).

where η_{1t} and Δ denote the white noise error term and the first difference operator, respectively. The parameters ς and λ are the short-run coefficients, and ζ_i, $i = 1, 2$ are the long-run coefficients of the ARDL model. The selection of the optimal lag is based on Schwarz – Bayesian information criterion (SBIC).[40] The bounds testing approach grounds on the joint F or Wald statistics, testing the significance of the lagged levels of the variables via the null hypothesis of no cointegration, $\mathrm{H_0}: \zeta_{1,2} = 0$ against the alternative of the existence of cointegration, $\mathrm{H_1}: \zeta_{1,2} \neq 0$. The asymptotic distribution of critical values is for cases in which the regressors are either purely I(0), purely I(1) or mutually cointegrated. The two sets of critical values are given in Pesaran et al. (2001), and its modified version for small samples, ranging from 30 to 80, is reported in Narayan (2005). Here, the critical values of Narayan (2005) are utilized for the bounds F-statistics due to the limited data availability of FLFPR and FSE.

The F-test has a non-standard distribution that depends on (i) the number of regressors (k), (ii) the sample size (n), (iii) whether the variables in the model are I(0) or I(1), and (iv) whether the model includes an intercept and/or a trend (Narayan, 2005). The upper value assumes that variables are I(1), whereas the lower value assumes that variables are I(0) in nature (Pesaran et al., 2001). If the F-statistics is above the upper bound, the null hypothesis is rejected at the significance level of the concerning bound, implying the existence of cointegration between the dependent variable and the regressor. However, if the F-statistics is below the lower bound, the null hypothesis fails to be rejected. The cointegration test is inconclusive if the computed F-statistics falls between the critical values. Table 5.3 shows the estimated ARDL models and the bounds F-test for cointegration results for both indicators of the labor market gender equality for the seven countries.

[40] Pesaran and Shin (1999) state that SBIC is more consistent than Akaike information criterion (AIC) and Hannan – Quinn information criterion (HQ). In addition, Monte Carlo evidence shows that SBIC and AIC determines reliable lag order (Panopoulou and Pittis, 2004; Emran, Shilpi and Alam, 2007).

TABLE 5.3: Estimated ARDL Models and Bounds F-test for Cointegration Results

	$\mathrm{F}(FLFPR\|y)$			$\mathrm{F}(FSE\|y)$		
Country	Optimal Lag	ARDL Model	F-statistics	Optimal Lag	ARDL Model	F-statistics
Canada	2	2,0	11.126***	2	1,1	7.291***
France	1	1,1	2.635	1	1,1	7.889***
Germany	1	1,0	5.131**	1	1,0	5.188**
Italy	1	1,0	0.234	1	1,0	0.417
Japan	1	1,0	0.542	1	1,0	0.711
UK	1	1,0	5.341**	2	1,2	6.194**
USA	1	1,0	4.867**	1	1,1	6.919***

Notes: The null hypothesis is the absence of cointegration between the dependent variable and the regressor. F-statistics are obtained from the ARDL cointegration test. The critical values from Narayan (2005, Appendix, Critical values for the bounds test: Case II: restricted intercept and no trend for k=1 and n=30) are 6.027 (1%), 4.090 (5%), 3.303 (10%) for the lower bound I(0); 6.760 (1%), 4.663 (5%), 3.797 (10%) for the upper bound I(1). The superscripts *** and ** denote the rejection of the null hypothesis at 1% and 5% levels, respectively.

The bounds F-test for cointegration analysis yields evidence of a long-run relationship between FLFPR and real GDP per capita at 1% significance level for Canada, and at 5% significance level for Germany, UK and USA. The long-run relationship between FSE and real GDP per capita is obtained at 1% significance level for Canada and USA, and at 5% significance level for Germany and UK. In addition, although France does not pass the bounds F-test for cointegration for FLFPR specification of gender equality in the labor market, it passes the test at 1% significance level for FSE specification, implying the existence of cointegration between FSE and real GDP per capita. However, the results for Italy and Japan imply no long-run relationship between gender equality in the labor market and real GDP per capita.

Subsequently, long-run and short-run models are estimated with Equations (5.12) and (5.13), respectively for those countries confirming the cointegration between FLFPR and/or FSE and real GDP per capita.

$$GE_t = \varphi_2 + \sum_{m=1}^{n} \varsigma_{2m}\, GE_{t-m} + \sum_{p=0}^{r} \lambda_{2p}\, y_{t-p} + \tau_2 D_{1t} + \omega_2 D_{2t} + \eta_{2t} \quad (5.12)$$

$$\Delta GE_t = \varphi_3 + \sum_{m=1}^{n} \varsigma_{3m}\, \Delta\, GE_{t-m} + \sum_{p=0}^{r} \lambda_{3p}\, \Delta y_{t-p} + \mu\, ECT_{t-1} + \tau_3 D_{1t} + \omega_3 D_{2t} + \eta_{3t} \quad (5.13)$$

where ECT denotes error-correction term determining the speed of convergence of the variables to the equilibrium. ECT is defined in Equation (5.14).

$$ECT_t = GE_t - \varphi_2 - \sum_{m=1}^{n} \varsigma_{2m} GE_{t-m} - \sum_{p=0}^{r} \lambda_{2p} y_{t-p} - \tau_2 D_{1t} - \omega_2 D_{2t} \qquad (5.14)$$

μ is the coefficient of ECT_{t-1}, and it should be statistically significant and negative. The long-run and short-run model estimations for FLFPR and FSE specifications of the labor market gender equality are presented in Tables 5.4 5.5, respectively.

TABLE 5.4: Estimated Long-run and Short-run Coefficients for Income Elasticity of FLFPR, 1984-2014

Indicator of GE: FLFPR				
Country	Canada	Germany	UK	USA
$T_{B,1}$	1989	1989	1989	1999
$T_{B,2}$	1995	1991	2001	2003
Estimated long-run coefficients				
Dependent Variable: GE				
y	0.331(8.38)***	0.666(13.10)***	0.180(2.31)**	0.413(76.37)***
$Constant$	0.818(2.01)*	-2.791(-5.39)***	2.372(3.01)***	3.913(2.19)**
D_1	0.013(0.88)	-0.018(-0.90)	-0.006(-0.20)	-0.042(-1.41)
D_2	-0.010(-0.82)	0.045(2.42)**	-0.009(-0.41)	-0.039(-1.73)*
Estimated short-run coefficients				
Dependent Variable: ΔGE				
Δy	0.105(3.00)***	0.370(4.03)***	0.044(1.62)	0.005(0.19)
$\Delta GE(-1)$	0.333(2.39)**	-	-	-
$\Delta Constant$	0.258(2.05)**	-1.551(-3.11)***	0.584(1.47)	0.514(3.30)***
ΔD_1	0.004(0.76)	-0.010(-0.87)	-0.002(-0.22)	-0.006(-1.38)
ΔD_2	-0.003(-0.73)	0.025(2.21)**	-0.002(-0.43)	-0.005(-1.24)
$ECT(-1)$	-0.316(-3.59)***	-0.556(-4.63)***	-0.294(-3.49)***	-0.131(-2.46)**
Adjustment in years	3.16	1.8	3.4	7.64
ARDL estimates				
Model	2,0	1,0	1,0	1,0
Adjusted R^2	0.989	0.986	0.970	0.969
RSS	0.001	0.004	0.001	0.0005

Notes: "(-1)" refers one lag of the associated variable. t-statistics for coefficients are in parentheses. RSS is the residual sum of squares. The superscripts ***, ** and * denote the statistical significance at 1%, 5% and 10% levels respectively.

Table 5.4 reports the estimated long-run and short-run coefficients based on the associated ARDL models by using FLFPR as the indicator of gender equality in the labor market. The results imply that the estimated long-run income elasticity of FLFPR is positive and statistically significant at 1% level for Canada, Germany and USA, and at 5% level for UK. That is, a positive response of FLFPR to real GDP per capita is provided in all countries confirming long-run relationship between the two through the bounds F-test for cointegration. In the ECM, μ is the coefficient of $ECT(-1)$, and it shows the adjustment speed towards the equilibrium following a shock to the system. A higher μ in absolute value implies a faster adjustment process. The estimated μ is negative and statistically significant at 1% level for Canada, Germany and UK, at 5% level for USA. The ECM shows that, within the cointegration model, there is a correction of the disequilibrium conditions at the following speeds: 31.6% for Canada, 55.6% for Germany, 29.4% for UK, and 13.1% for USA. Hence, the speed of convergence to long-run equilibrium in years is approximately three for Canada and UK, two for Germany, and eight for USA.

Table 5.5 presents the estimated long-run and short-run coefficients based on the associated ARDL models by indicating the labor market gender equality with FSE. The results reveal that the estimated long-run income elasticity of FSE is positive at 1% statistical significance level for Canada, France, Germany, UK and USA. In other words, a positive relationship between FSE and real GDP per capita is satisfied in the long run in all countries passing the bounds F-test for cointegration. The estimated μ is negative and statistically significant at 1% level for Canada, Germany and UK, and at 5% level for France and USA. The ECM highlight that, within the cointegration model, there is a correction of the disequilibrium conditions at the following ratios: 31.7% for Canada, 18.5% for France, 37% for Germany, 38.1% for UK, and 17.7% for USA. In other words, the speed of adjustment towards long-run equilibrium is approximately three years for Canada, Germany and UK, five years for France, and six years for USA.

TABLE 5.5: Estimated Long-run and Short-run Coefficients for Income Elasticity of FSE, 1984-2014

			Indicator of *GE: FSE*		
Country	Canada	France	Germany	UK	USA
$T_{B,1}$	1991	1989	1989	1989	2006
$T_{B,2}$	2009	2001	1991	1992	2009
			Estimated long-run coefficients		
			Dependent Variable: GE		
y	0.145(7.68)***	0.339(2.68)***	0.385(8.63)***	0.352(62.31)***	0.358(124.49)***
$Constant$	2.322(12.02)***	0.350(0.27)	-0.243(-0.53)	3.509(12.88)***	3.430(5.25)***
D_1	0.029(4.36)***	-0.016(-0.67)	-0.004(-0.22)	0.010(1.14)	-0.008(-0.57)
D_2	0.004(0.66)	-0.007(-0.28)	0.027(1.80)*	0.041(5.62)***	0.010(0.83)

TABLE 5.5: Continued

			Estimated short-run coefficients		
			Dependent Variable: ΔGE		
Δy	-0.143(-4.56)***	-0.141(-2.11)**	0.142(2.82)***	-0.110(-2.19)**	-0.115(-2.40)**
$\Delta y(-1)$	-	-	-	-0.097(-2.37)**	-
$\Delta Constant$	0.737(3.44)***	0.065(0.28)	-0.090(-0.49)	1.337(4.93)***	0.718(2.74)***
ΔD_1	0.009(2.18)**	-0.003(-0.63)	-0.001(-0.21)	0.004(1.05)	-0.002(-0.50)
ΔD_2	0.001(0.60)	-0.001(-0.27)	0.010(1.69)*	0.016(3.20)***	0.002(0.70)
$ECT(-1)$	-0.317(-3.52)***	-0.185(-2.42)**	-0.370(-3.52)***	-0.381(-3.85)***	-0.177(-2.04)**
Adjustment in years	3.15	5.41	2.7	2.62	5.65
			ARDL estimates		
Model	1,1	1,1	1,0	1,2	1,1
Adjusted R^2	0.993	0.989	0.989	0.988	0.962
RSS	0.0002	0.001	0.001	0.0003	0.0003

Notes: "(-1)" refers one lag of the associated variable. t-statistics for coefficients are in parentheses. RSS is the residual sum of squares. The superscripts ***, ** and * denote the statistical significance at 1%, 5% and 10% levels respectively.

Table 5.6 demonstrates an interesting coincidence between business cycles dates (BCs) declared, and the structural break dates on the labor market gender equality series ($T_{B,1}$ and $T_{B,2}$), identified by NP unit root tests.[41] First, $T_{B,2}$ of FSE coincide perfectly with the BCs for all countries but France. Second, $T_{B,1}$ of FSE is identical for France, Germany and UK; $T_{B,1}$ of FLFPR is identical for Germany and UK; $T_{B,1}$ of both FLFPR and FSE for Canada show very similar pattern with that for Continental Europe countries. On the other hand, $T_{B,1}$ of both FLFPR and FSE for USA follow a different pattern.

TABLE 5.6: Structural Break Dates versus Business Cycle Dates

		FLFPR				*FSE*		
Country	$T_{B,1}$	BCs	$T_{B,2}$	BCs	$T_{B,1}$	BCs	$T_{B,2}$	BCs
Canada	1989	Mar90-Mar92	1995	na	1991	Mar90-Mar92	2009	**Jan08-Jul09**
France	-	-	-	-	1989	Feb 92-Aug 93	2001	Aug 02-May 03
Germany	1989	Jan91-Apr94	1991	Jan91-Apr94	1989	Jan91-Apr94	1991	**Jan91-Apr94**
UK	1989	May90-Mar92	2001	na	1989	May90-Mar92	1992	**May90-Mar92**
USA	1999	Mar01-Nov01	2003	na	2006	Dec07-Jun09	2009	**Dec07-Jun09**

na: BCs not observed in ±2 years

[41] The BCs are from Economic Cycle Research Institute (ECRI). $T_{B,1}$ and $T_{B,2}$ refer to the first and the second structural break years of the labor market gender equality series, respectively. When including the BCs, ±2 of $T_{B,1}$ and $T_{B,2}$ are considered.

The NP unit root test results show that all series are exposed to structural breaks for all the seven countries. To this end, the stability of the long-run and short-run coefficients should be detected through the cumulative sum (CUSUM) and cumulative sum of squares (CUSUMSQ) tests due to Brown, Durbin and Evans (1975). Figures 5.1 and 5.2 illustrate the plots of CUSUM and CUSUMSQ test statistics falling inside the critical bounds of 5% significance level, which indicates that the estimated parameters for both FLFPR and FSE specification of gender equality are stable over the periods. To sum up, gender equality in the labor market shows a positive trend in response to income in Canada, France, Germany, UK and USA in the period 1984-2014. In addition, the statistically significant and negative coefficients of $ECT(-1)$ indicate that any deviation from the long-run equilibrium is corrected to return to the long-run equilibrium level for each period.

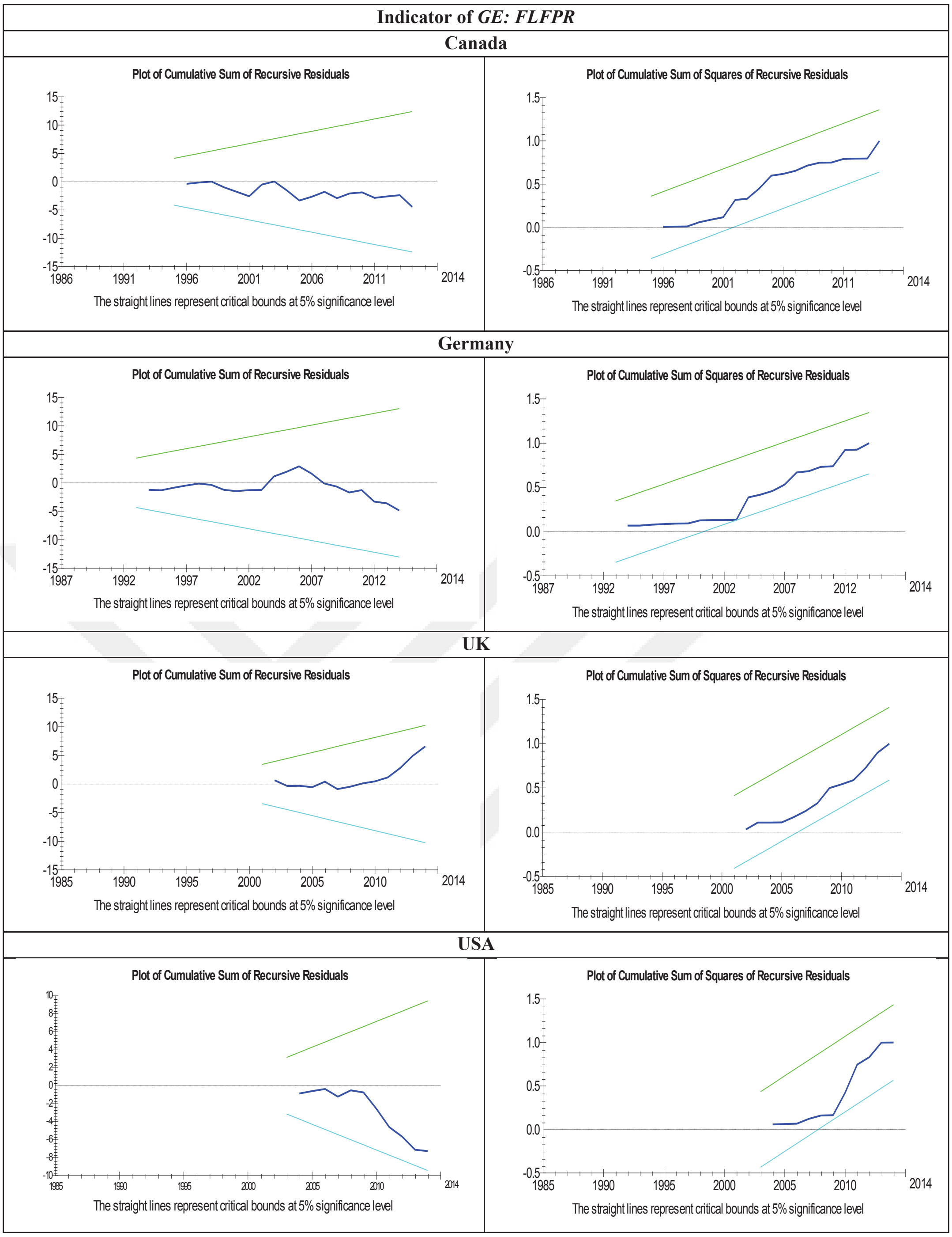
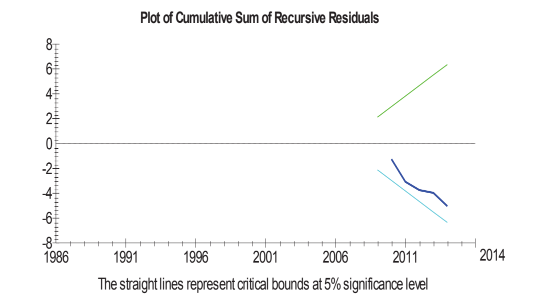
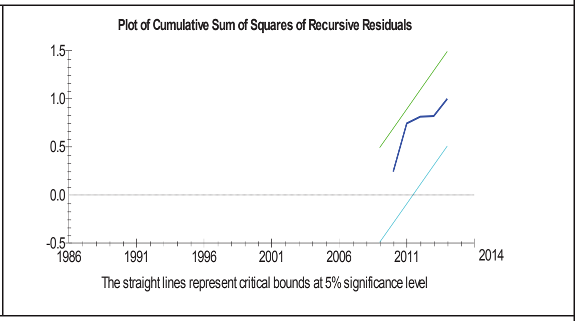


FIGURE 5.1: Plot of CUSUM and CUSUMSQ Tests for parameter stability (FLFPR)

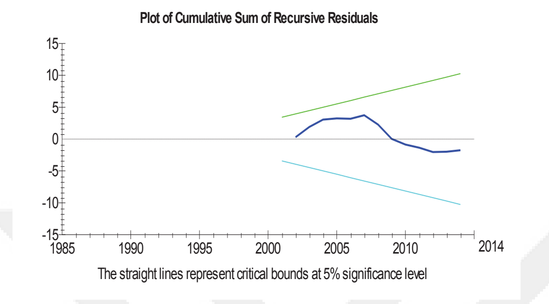
Indicator of *GE: FSE*

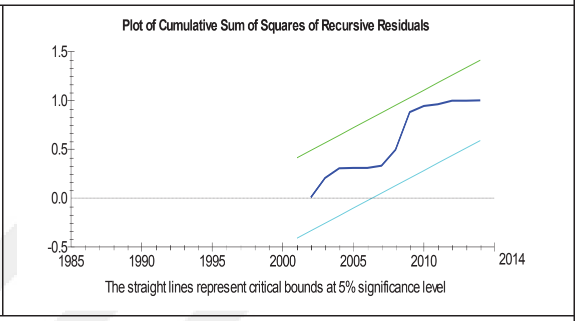
Canada



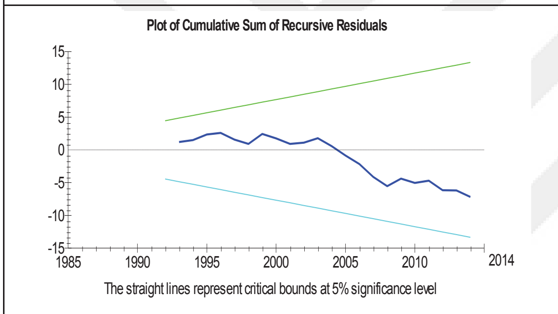


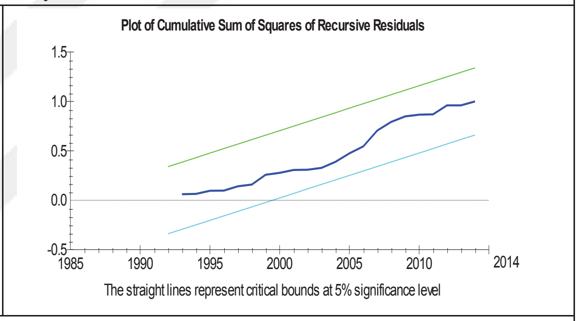
France



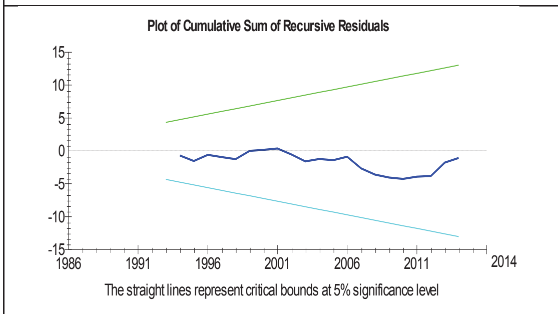


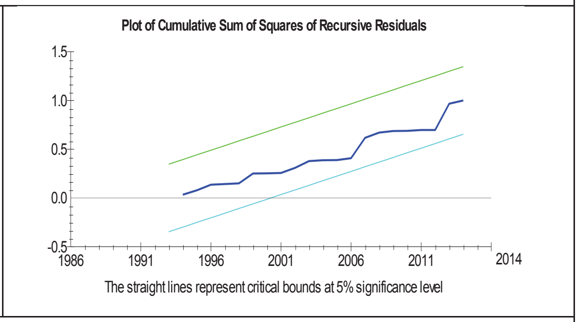
Germany



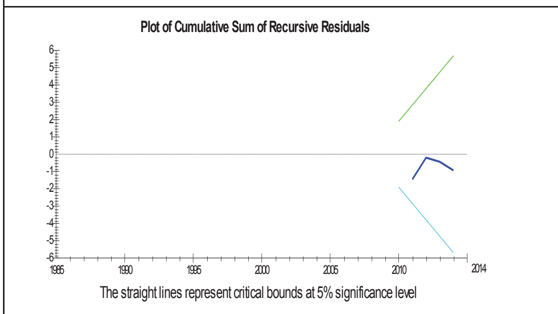


UK





USA



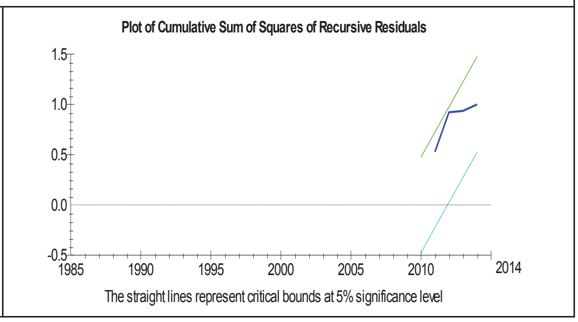


FIGURE 5.2: Plot of CUSUM and CUSUMQ Tests for parameter stability (FSE)

5.4 Concluding Remarks and Discussions

The contribution of this chapter to the literature is twofold. The first is to investigate whether the impacts of shocks on FLFPR and FSE, the two potential indicators of gender equality in the labor market, are permanent or transitory in nature, when structural breaks are allowed for in the process. The second is to determine long-run and short-run income elasticities of FLFPR and FSE, by additionally taking into account identified structural break dates of these series. The empirical analyses cover G7 countries over the period 1984-2014. These incorporated analyses enlighten policy makers about the degree of resistance of the labor market gender equality series to shocks, especially during economic downturns. To this end, the stationarity features of FLFPR and FSE series are analyzed at first through NP (2010) unit root test, which allows for endogenously determined structural breaks at two time points. The test results reveal that FLFPR and FSE series are non-stationary processes, which imply that the impacts of shocks on gender equality in the labor market are permanent for all G7 countries in the period 1984-2014. Subsequently, the existence of cointegration between income and gender equality in the labor market is detected and long-run and short-run models are estimated through ARDL bounds testing approach of cointegration. The estimates reveal that real GDP per capita has a positive impact on FLFPR in Canada, Germany, UK and USA, on FSE in also France during the sample period. In all, this chapter concludes that the labor market gender equality series are non-stationary processes, and that they show positive response to income in five out of G7 countries. Hence, any direct or indirect (via income) negative shock on these series may lead to permanent breaks on gains in gender equality in the labor market. To counteract this tendency, during downfall periods, policy makers should develop stabilization policies and incentive mechanisms aimed at the resilience of women’s participation in the labor market, and allowing gender equality to be sustained.

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PUBLICATIONS

Accepted to be published in journals covered by SCI, SSCI or AHCI

Biresselioglu, M. E., D. Kılınç, E. Onater, and T. Yelkenci. 2016. *Estimating the Political, Economic and Environmental Factors' Impact on the Installed Wind Capacity Development: A System GMM Approach*. Accepted to be published in Renewable Energy (SCI), Vol. 96 (Part A): 636-644.

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SELECTED CONFERENCE PRESENTATIONS

First Annual All-Izmir Economics Workshop, May 2016, Yaşar University, Izmir, Turkey. Paper presented: *Convergence in Financial Development Measures across the EU-15*. Web: http://iktisat.yasar.edu.tr/en/category/activities/.

2nd World Keynes Conference, September 2015, Pamukkale University, Denizli, Turkey. Paper presented: *Heterogeneous 'Gender Kuznets Curve'*. Web: http://keynes.pau.edu.tr/.

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