Unlocking the Gender-Economic Growth Nexus in Developed, Semi-Industrialized, and Low-Income Agricultural Economies

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ABSTRACT

This research examines the intricate relationship between gender equality and economic growth in developed, semi-industrialized, and low-income agricultural economies. By assessing the impact of gender wage inequality on economic growth using a panel dataset of 46 countries, we uncover distinct patterns influenced by economic structure and development stage.

Our findings reveal that the short-term effect of the gender wage gap differs across economies, acting as a stimulus to growth in semi-industrialized nations but reducing demand and growth in the long run. Conversely, low-income agricultural economies and developed countries experience negative short-term effects of gender wage inequality but may benefit from long-term growth prospects. This study contributes to the feminist macro literature by providing empirical evidence of the role of gender wage inequality on economic growth while accounting for country-specific variations in production structure and development stage.

Our results highlight the importance of gender equality in shaping aggregate demand and its implications for sustainable growth.

KEYWORDS: Gender equality; demand-led growth; economic development; autoregressive distributed lag models (ARDL).

I. INTRODUCTION

A revived interest in the relationship between income distribution and the macroeconomy is evident in the academic literature and amongst policy makers (Stockhammer 2011, 2013; Stockhammer and Wildauer 2015; Kim, Lima, and Setterfield 2019). The sustained deceleration in global economic growth, the global financial crisis of 2007-09, and the economic effects of the 2020 global pandemic have renewed interest in this strand of economic research, with emphasis on a reevaluation of the assumed economy-wide benefits of mainstream policy recommendations that favour wage moderation and austerity (Rajan 2010; Wisman and Baker 2011; Stiglitz 2010; Wade 2011; 2011; Hein et al. 2012; Wisman 2013; Mellish et al. 2020; Nikoforos 2020).

Insights from this large body of research on the macroeconomic effects of inequality suggest that gender inequality may affect the rate of economic growth, and macro-level policies have been found to have gendered effects. Among neoclassical analyses, researchers find that gender gaps in education, life expectancy, and employment lower a nation's living standard and the rate of productivity growth (Hill and King 1995; Knowles et al. 2002; Klasen and Lamanna 2009). Most of these studies assume full employment and therefore emphasize long-run gender effects, ignoring short-run demand-side impacts of changes in gender equality. Further, these studies implicitly assume a role for gender inequality that is similar across countries. That is, their theoretical and empirical methodologies assume we can make universal claims about the effect of gender inequality on economic performance.

Heterodox analyses, however, introduce structural parameters into macro models to reflect differences in the structures of production as well as macro-level policies that influence output. Further, feminist macro models account for the effect of gender job segregation that varies with a country's structure of production (Braunstein 2008; Seguino 2019; Onaran 2016). This study adds to the existing feminist macro literature by econometrically estimating the role of gender wage inequality on economic growth, accounting for country variation in the structure of production and stage of development.

We begin by highlighting the key features of countries at different stages of development and structures of production. In analyzing the macroeconomic effect of gender wage inequality, we take account of the gender division of labor within the paid economy (Braunstein 2008; Seguino 2010a, 2010b; Onaran 2016). While gender job segregation is a global phenomenon, the extent to which it persists, and the nature of occurrence varies across economic structures. The distribution of jobs by gender, combined with women's lower wages on average, have crucial implications for gendered income and wage inequality (Mandel and Semyonov 2014). With this framing of the role of gender, we adopt a demand-led macroeconomic model, which highlights an under-consumptionist view of growth, to analyze the potential effects of income distribution and gender wage inequality on aggregate demand in the short run as well as long-run growth.

On the theoretical side, this study relies on a variation of the gendered post-Kaleckian demand-led growth model that has been widely used in recent years, whereby gender inequality affects the various components of aggregate demand. The Kaleckian approach, which emphasizes the role of monopoly power in determining the functional income distribution, is combined with a framework in which the distribution of wages between male and female workers is determined by relative bargaining power. Gender inequality(proxied by the gender wage ratio) is captured by women's share of the wage bill.

On the empirical side, our contribution is to shed light on possible heterogeneity of the effects across countries at various stages of development. We therefore report separate results for high-income, middle-income, and low-income countries or, alternatively following Seguino (2010a), advanced/developed economies (DC), semi-industrialized economies (SIEs), and low-income agricultural economics (LIAEs). By modeling growth as a function of women's share of the wage bill, we are able to identify empirically countries that are gender cooperative (a higher female share stimulates growth) as compared to gender conflictive (a restriction to women workers slows growth). Our model uses panel data on a sample of 31 countries over the period 1970-2011. The panel data approach that considers cross-country heterogeneity, dynamics, and the possibility of cross-sectional dependence is employed.

Overall, our findings suggest that global economic growth is wage-led in the long run and profit-led in the short run. Following our investigation, we find that economic growth is gender cooperative in the long run but gender conflictive in the short run. Also, given that we find short-run contractionary effects of gender inequality on growth globally (and for the developed and middle-income countries), two points may be considered. First, wage-led growth does not always imply gender equality-led growth, and vice versa. Finally, while our initial gender effects provide us with information about the growth trajectory of an economy, our later estimations provide us with some information about the growth contributions of gender equality to the relative components of aggregate demand.

II. LITERATURE AND THEORETICAL REVIEW

1.1. GENDER AND ECONOMIC GROWTH

1.1.1. Demand - Led Growth

Kalecki (1971, 1989) formulated a theory of national income determination established on the principle of demand-led growth with a potential for excess capacity. Another strand of the literature on distributional effects on aggregate demand follows from the Goodwin (1967) business cycle model. The Goodwin model also builds on Marxist principles but, unlike Kaleckian models, assumes that investment is saving-determined and that profits are the main determinants of saving; as such, the Goodwin cycle attributes lower unemployment to increases in the profit share. Another distinction between both approaches is the tendency for empirical analyses following the Goodwin (1967) model to employ a systems-based approach (usually VAR) in their estimations (Stockhammer and Onaran 2004; Barbosa-Filho and Talyor 2006; Flaschel and Proano 2007; Allain and Canry 2008). The post-Kaleckian models on the other hand typically employ a single equations approach, estimating separate functions for the different components of aggregate demand (Bowles and Boyer 1995; Naastepad and Storm 2006; Hein and Vogel 2008; Stockhammer, Onaran, and Ederer 2008; and Stockhammer and Wildauer 2016).

Both approaches have their strengths and shortcomings. For the former, the systems approach allows for endogeneity of the distribution variables and as such takes simultaneity into account. However, this approach often suffers from omitted variable bias and only provides a weak measure for identifying distributional effects on consumption and investment. The single equation approach, on the other hand, is good at identifying these distributional effects but fails when it comes to dealing with problems posed by endogenous variables in the regression model.[1]

In this study, we follow the Keynesian (Keynes, 1936, p.14), Kaleckian (Kalecki, 1971) and Davidsonian (Davidson, 1972, pp. 344 and 349) postulations that wage bargaining takes on significance only when we explicitly take into account heterogeneity of wages across groups of workers and capitalists. An essential feature of this post-Kaleckian growth model is the delineation between wage-led and profit-led growth, mostly credited to the works of Blecker (1989) Bhaduri and Marglin (1990) and Taylor (1991). Studies have found that smaller and/or more open economies tend to be profit-led, whereby a redistribution from workers to capitalists is a stimulus to output, employment, and growth (Fernandez 2005; Barbosa-Filho and Taylor 2006; Storm and Naastepad 2012; Onaran and Galanis 2012; Keifer and Rada, 2014; Carvalho and Rezai 2015). This is in part attributable to the effect of a redistribution on exports and imports, although macro-level policies also play a role. In contrast, more closed economies tend to be wage-led (Onaran and Galanis, 2012; Blecker xxx). A common feature of these findings is that they fail to distinguish between long- and short-run aggregate demand effects and do not account for the role of gender on these macroeconomic aggregates. The goal of this paper is to examine the multifaceted linkages between gender wage inequality, development, and demand-led growth.

1.1.2. Gender Equality-Driven, Demand-led Growth and Economic Structure: Some Stylized Facts

The Kaleckian demand-led growth model assumes, *ceteris paribus*, that redistributing income from the rich to the poor will lead to a higher rate of growth, especially for an economy under a wage-led growth regime.[2] In a similar vein, it would be expected that redistributive policies that promote an upward convergence of females’ wages to males’ will also contribute positively to economic growth. However, this hypothesized relationship may differ for economies at different stages of development. In an influential study using cross-country research, Seguino (2000) identified a negative relationship between gender equality in wages and economic growth for a set of semi-industrialised countries using data from 1975 to 1995. The findings suggest that higher levels of gender inequality in wages served to promote

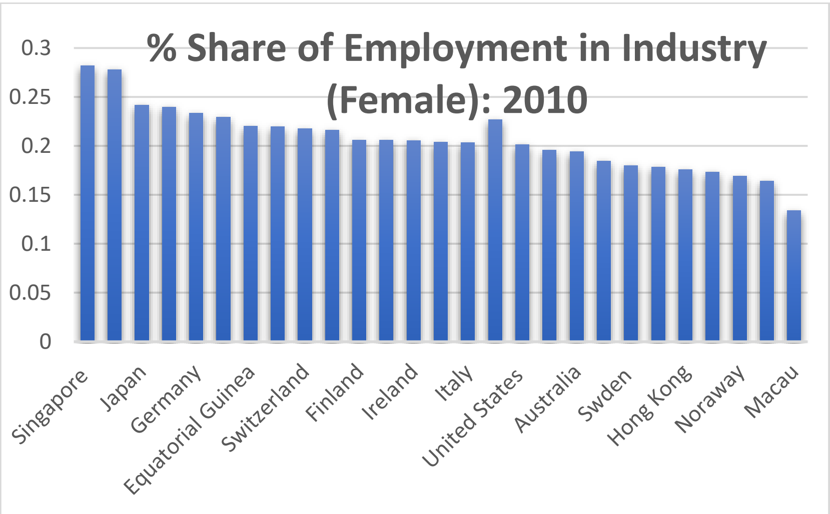
[1] This paper, unlike previous research using the single equations approach, attends to some of the endogeneity concerns through the use of procedural empirical techniques.
[2] This hypothesis is inferred because the proportion of consumption out of capitalist income is lower than the consumption out of workers’ income.

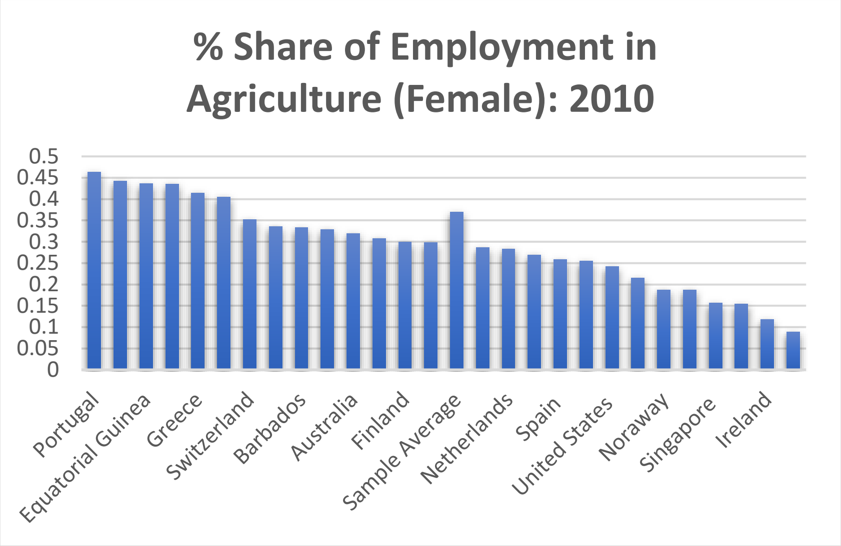
economic growth through a reduction in labour costs and a consequent stimulation of export demand. While there is still little evidence for developed economies, Badru (2018) finds for Canada, the UK and Australia that a more equitable wage distribution between men and women raises aggregate consumption due to the likelihood of a higher marginal propensity to consume (mpc) for women compared to men. This finding, in turn, may imply that eliminating gender wage gaps may foster demand-led growth.

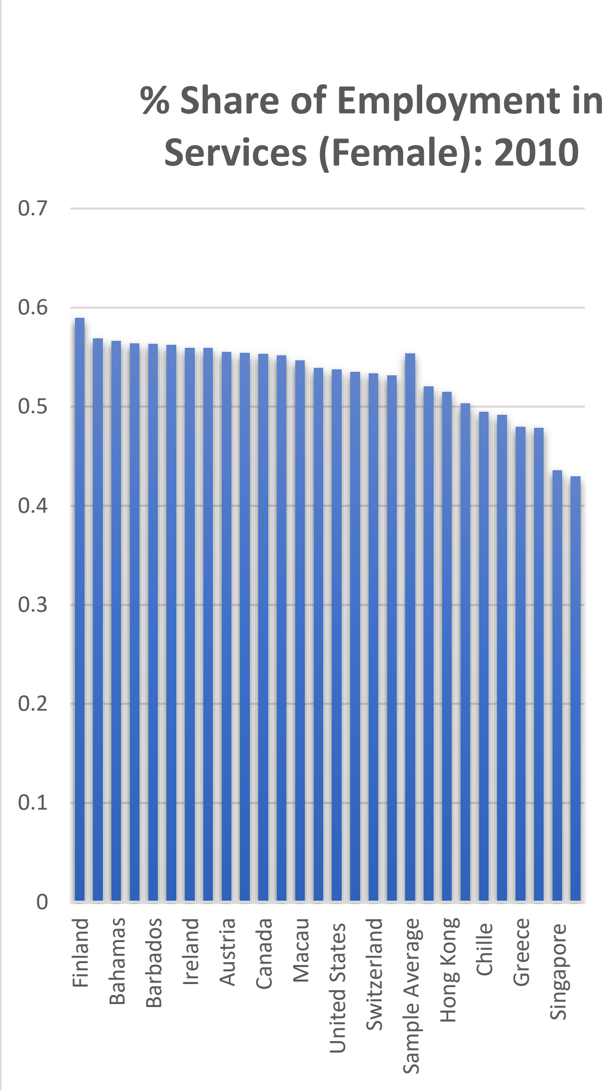
Gendered labour market outcomes also differ between country groups. In semi-industrialized export-oriented economies (SIEs), women workers are concentrated in the export sector, which produces labor-intensive manufactured goods, services, and non-traditional agricultural exports. In these regions, according to Seguino (2007), firms are often mobile, display monopsonist buyer of labor characteristics and can easily trade globally through exports or export platforms – especially selling a bulk of their products in developed regions – leading to increased mark-ups and greater comparative advantage (Busse and Spielmann, 2003). This implies that higher female wages can serve to dampen both investment and exports, producing an economic contraction and worsening the balance of payments (Bamber and Staritz, 2016; Seguino, 1997; 2007). In these countries, even if women's mpc is higher than, the expansionary effect of higher female wages is unlikely to be large enough to offset the negative investment and export effects. Seguino (1997, 2007) explains that this phenomenon may be driven by the mobile nature of labor-intensive firms which are prevalent in SIEs and which employ a large proportion of female workers in industries where the export demand for the goods women produce is price-elastic.

The reverse may be the case for the effect of a more gender-equal wage setting in low-income agricultural economies (LIAEs) – albeit through different transmission channels. First, women in LIAEs (whose jobs are often concentrated in informal sectors and in domestic work) spend a higher proportion of their income on childcare and home production (Pahl, 2000; Gummerson and Schneider, 2013). Another salient feature of low-income countries is that a lower proportion of workers are often concentrated in formal employment in comparison to higher income countries. More so, women dominate the unpaid labour force in home production and in the agricultural sector when compared to middle- and high-income countries (where women still bear a disproportionate share of such tasks) Data evidence suggests a high concentration of female employment in the services sector in high-income developed economies (ILO 2020).

According to Anker (1998), such concentration in specific occupations and sectors is often a result of a high supply of female labour in these industries, which could in turn result in a prevalence of lower wages for women as well as an increased level of unemployment amongst women. Standing (1989, 1999) posits that this segregation of women into particular sectors could result in the ‘feminisation’ of jobs in these industries; these jobs are often depicted as low skilled jobs with high turnover of employees and low associated bonuses. In the case where economies are markedly characterised with openness to the global economy, and as such largely export-oriented, gender-wage gaps are expected to persist and likely increase over time where such structures persist (Zveglich and Rodgers 2004; Berik 2008; Tran 2019). Hence, larger gender wage gaps in semi-industrialised economies have, in effect, influenced growth positively in the short-run due to the largely female workforce. Therefore, increased female wages may result in a loss in profits due to a reduction in global competitiveness. However, this result may vary in low-income agro-based economies where lower gender wage gaps may spur economic growth (Seguino 2006).



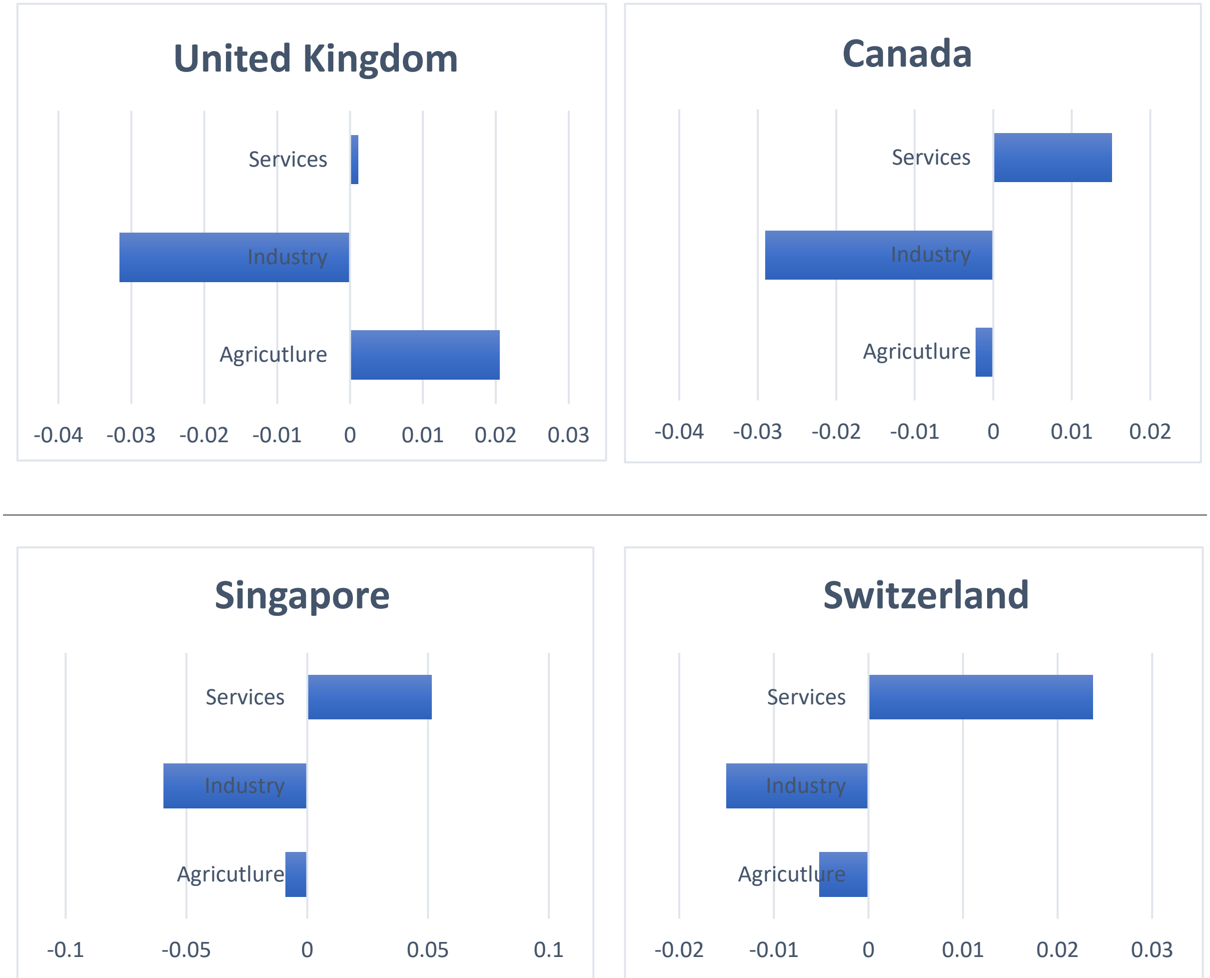


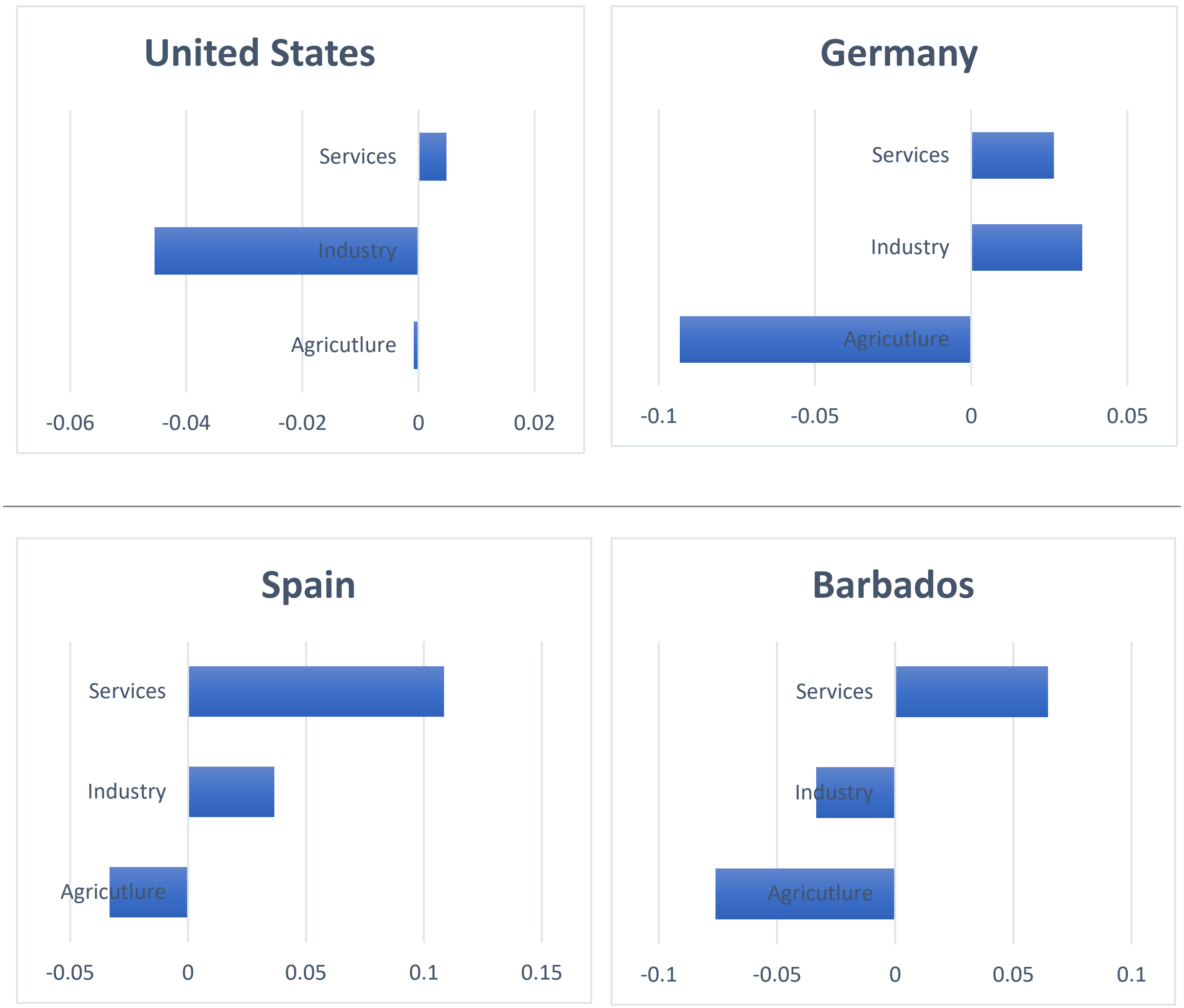


Source: International Labour Organization Data (ILO; 2021)

For Spain and Germany, the data (ILO, 2021) also suggests that the increases in female employment in services and industry respectively is mostly due to a larger entrant of [especially migrant] women into these sectors and less as a consequence of change in employment preferences. Ortiz-Ospina (2018) shows that higher female labour force participation in many high-income countries is often associated with fewer hours of work on average. This is likely to be due to in part to lack of affordable care – where women take on a larger proportion of unpaid caring work – and a need to maintain some work-life balance in two-parent households (Verick, 2014). Verick (2014) further explains a potential u-shaped relationship between female labour force participation and economic development as economic activity transitions from agriculture (LIAEs) to industry (SIEs) and then to services (DCs – especially for women). The transition from female work in industry to services can be attributed to higher levels of education, lower fertility rates and weaker societal gender norms which encourage women's labour market engagement. Given that such norms are not eliminated, and women still bear a disproportionate share of the unpaid care labour burden, a larger proportion of women move into the services sector which often allow for more work-time flexibility. In developing countries, high female labor force participation rates typically reflect poverty. Women earn less than men and are more likely to be engaged in unprotected jobs, such as domestic work. Education raises the reservation wage and expectations of women, but it needs to be matched by job creation.

Percentage Change in Female Sectorial Employment from 1991 to 2015





Using data collected on full-time US workers, Blau and Kahn (2007) observe that the female-to-male earnings ratio was approximately 0.60 during the period 1955 till the 1970s. This proportion increased by 10% in the 1980s and has since then only minimally improved; women's wages relative to men has improved by less than 5% since the 1980s. In a later study, Corbett and Hill (2012) observed a widening gap in gender wage inequality since the beginning of the 21st century, especially among highly skilled workers; this is in spite of a reduction in the overall wage gap. More so, Aisenbrey and Bruckner (2008) and Magnusson (2015) observe that, other than the prevalence of gendered occupational segregation and wage gaps, another common labour market phenomenon is within-occupation gender segregation and wage inequality.[3]

[3] Magnusson (2015) also finds that the gender wage gap that currently exists in the field of medicine – in Sweden – was considerably higher in 2007 than it was in 1975.

Several authors have suggested that this wage gap persists due to the weaker collective bargaining position of women in the labour market (Addison et al., 2017; Dustmann and Schönberg, 2009). Others have attributed the stagnation in the elimination of gender wage gaps to women's lower human capital position and changing labour market institutions (DiNardo et al., 1996, Borjas, 2002). Even so, we still can observe that gender wage gaps have not just persisted over time, they appear to be a global occurrence – irrespective of the level of economic development, the labour market structure, legal sanctions and economic reforms in nations across the globe. Female workers also appear to be more concentrated in certain sectors and industries (BLS 2011; DOL 2011; Ortiz-Ospina and Tzvetkova 2017).

Since the early 1980s, research has examined the links between gender wage equality and macroeconomic aggregates, such as trade, investment, and economic growth, and has established that macroeconomic policies have gendered effects. A valid standpoint is also to view the saving and consumption decisions of women as having a distinct effect on demand and hence growth, with reference to a wage-led regime. This is especially relevant from a demand-led growth perspective where increased consumption is expected to positively affect aggregate demand. It is useful to note that while this study employs a structuralist approach to understanding the macroeconomic implications of gender inequality following an underconsumptionist framework of analysis, there are several other notable theoretical approaches to modelling the relationship between gender and the macroeconomy. Seguino (2020) summarizes the main tenets of these approaches including both long-run neoclassical models, overlapping generations (OLG) model as well as short-run demand models that incorporate gender into the heart of the economic analysis.

1.2. THEORETICAL MODEL

Demand-led growth models in the heterodox/Keynesian tradition have long emphasized the importance of income distribution and aggregate demand with relevance to wages as a source of demand to stimulate growth. Kaleckian macroeconomic models assume a significant role for the aggregate level of spending, which is a function of the distribution of income between workers and capitalists. This is a structural rather than an individualist disaggregation based on the different economic functions of workers and capitalists and corresponds to the institutional division between firms and households. This is a framework from which the employment of gender as an organizing framework in macroeconomics can be effectively introduced, as any such disaggregation ought to be based on a similar understanding of the way in which gender as a social institution impinges on or constrains the behavior of the macro economy.

However, the effect of gender inequality on economic performance may vary for different economies and such varied outcomes may depend on the conceptual period being investigated (i.e. short-run or long-run), the mains drivers of productive activity, the stage of economic development and the prevailing institutional structure of the economy (Blecker and Seguino, 2002; Seguino, 2012; Doepke and Tertilt, 2019). To model this potentially bidirectional relationship between gender and the macroeconomy, we consider a structuralist analytical approach which allows for some institutional and macroeconomic heterogeneity – in particular, we explore this relationship for developed economies (DCs), semi-industrialized economies (SIEs) and low-income agricultural economies (LIAEs). Countries are divided into these structural groups based on their level of industrial intensity index (III) and financial development.[4]

1.2.1. Gender and Growth

Gender macro models build from the structuralist post-Keynesian/post-Kaleckian macroeconomic models discussed above which, although originally gender-blind, differ from neoclassical growth models in that they pay particular attention to the demand side of the economy in the determination of growth and output (Blecker, 1989; Bowles and Boyer, 1995; Marglin and Bhaduri, 1990; Taylor, 1991). Feminist economists have adapted this framework to account for gender differences in wages, thereby simultaneously exploring the effects of both inter-class and intra-class distribution (Grown et al., 2002; Braunstein, 2012; Seguino 2020).

We extend Bhaduri and Marglin's (1990) model of the original Keynes/Kalecki framework by incorporating the impact of the gendered distribution of wages on aggregate demand. First, we start by defining new variables to represent the wages of our gender groups such that $W_m(t)$ and $W_f(t)$ represent the wages of men and women at time *t*, respectively. Secondly, we extend the wage share specification (*WS*) in the Bhaduri and Marglin model to

[4] To do this we weigh the III for each country (using 2018 wdi data) by the IMF financialization development data. Countries above 50% are DCs, countries within 40 – 60% are SIAEs and countries less than 40% are LIAEs. See appendix for more information on how this measure is computed. Our final sample of countries is represented in appendix 4.

incorporate a gendered wage distribution such that $W_m(t) + W_f(t)/PY$ is our wage share, where P is the price level, and Y is output. To incorporate the gender wage gap, we introduce a new variable, $GG(t)$, which represents the ratio of female to male wages $\frac{Wf(t)}{Wm(t)}$. This allows us to model the impact of changes in the gender wage gap on aggregate demand (AD).

We allow wages to be weakly exogenous, firstly because we account for the role of effective demand (and as such unit labor costs/wage rates) in the determination of output while understanding that this co-determination of wages with the real level of output can be mitigated by policies allowing for fluctuations in the exchange rate in the case of an open economy – making changes in the real wage neither strictly endogenous or exogenous. This is especially true in open economies due to participation in international trade and, as such, global price fluctuations and a push for competitiveness.

1.2.2. Aggregate Demand Components

To examine the nexus between income distribution and aggregate demand, we propose a framework based on the Bhaduri and Marglin (1990) formulation of the Post-Keynesian demand-led growth model (Keynes 1936; Kalecki 1969; 1971). The adopted theoretical model for this analysis closely follows the approach by Stockhammer et al. (2008), and adds to this the gendered impact of income inequality within a panel framework.

Assuming an open economy with no government intervention, production can be assumed to occur through three different streams: stream 1 produces investment goods (*I*); stream 2 produces goods for capitalists and workers (*C*) and the third produces goods for foreign trade (*NX*). The public sector is captured by *G* which is treated as exogenous. The behavioural functions for aggregate demand, incorporating the gendered distribution of wages, can then be stated as:

$$Y = C\left(W_m(t), W_f(t), GG(t)\right) + I + NX + G \tag{1}$$

where the consumption function C is now a function of the wages of men and women, and the gender wage gap. The previously defined gendered variables can then be incorporated into the following behavioral equations:

$$C = f(Y, WS, DH, LP, GG) \tag{2}$$

$$I = f(Y, WS, INT, DH, DB, GG) \tag{3}$$

$$NX = f(Y, WS, FY, EX, LP, GG) \tag{4}$$

$$G = G' \tag{5}$$

where *Y* and *WS* are real disposable income, the ratio of female to male wages and the wage share respectively. *DH* and *DB* are the household debt-to-GDP ratio and the business debt to *GDP* ratio. *INT, FY, EX* and *LP* represents the long-term interest rate, foreign income in real terms, the nominal effective exchange rate and female labor force participation, respectively.

Due to our focus on the private sector in this study, government expenditure (*G*) is treated as exogenous. In the following sub-sections, we provide the rationale for incorporating the right-hand side variables in the specification of the above stated behavioral equations.

Consumption

Consumption is a component of aggregate demand and is assumed to be dependent on income such that total income (*Y*) in the economy is derived from two sources: capital income (Π) and labor income (*W*); with labor income consisting of wages of men (W_m) and women $\left(W_f\right)$. We can write this as:

$$Y = \Pi + W_m + W_f \tag{6}$$

Differentiating with respect to income Y, we get the shares of profits (s_Π) and wages (s_W) in income:

$$s_\Pi = \frac{\partial \Pi}{\partial Y} = \frac{\Pi}{Y} \tag{7}$$

$$s_W = \frac{\partial (W_m + W_f)}{\partial Y} = \frac{W}{Y} \tag{8}$$

such that: $s_\Pi + s_W = 1$

Assuming a non-gendered delineation for profits, we only define the share of female wages (s_Wf) and male wages (s_Wm) in total wages as:

$$s_{Wf} = \frac{W_f}{\left(W_m + W_f\right)} \tag{9}$$

$$s_{Wm} = \frac{W_m}{\left(W_m + W_f\right)} \tag{10}$$

Let us assume that that the share of profits in total income increases as total income increases while the share of wages in total income decreases as total income increases, holding everything else constant. Differentiating (9) and (10) with respect to Y, we get $\frac{ds_{\Pi}}{dY} = -\frac{\Pi}{Y^2}$ and $\frac{ds_W}{dY} = -\frac{W_m + W_f}{Y^2}$ indicating that: If $\frac{d\Pi}{dY} > 0$ (i.e., profits increase with total income), and If $\frac{\Pi}{Y} > 0$ (i.e., profits constitute a positive share of total income), Then $\frac{ds_{\Pi}}{dY} > 0$, while If $\frac{dW_m}{dY} + \frac{dW_f}{dY} < 0$ (i.e., wages decrease with total income), and If $\frac{W_m + W_f}{Y} > 0$ (i.e., wages constitute a positive share of total income), then $\frac{\mathrm{ds_W}}{\mathrm{dY}} < 0$.

The income of each individual i (female and male) can be defined as the sum of their wage income and their share of profits:

$$Y_i = s_{\Pi} * \Pi + s_W * W_i \tag{11}$$

The sum of all individual incomes equals total income, such that $\Sigma Y_i = Y$.

Furthermore, we assume that there is a higher propensity to save on profits $(\mathrm{s_{\Pi}})$ than on wages $(\mathrm{s_W})$, such that $\mathrm{s_{\Pi}} > \mathrm{s_W}$. Similarly, we assume that the propensity to consume from wages $(\mathrm{c_W} = 1 - \mathrm{s_W})$ is higher than the propensity to consume from profits $(\mathrm{c_{\Pi}} = 1 - \mathrm{s_{\Pi}})$, so we have $\mathrm{c_W} > \mathrm{c_{\Pi}}$.

Given that average wages in most sectors are significantly larger for male workers, the saving rate $(\mathrm{s_i})$ and marginal propensity to consume $(\mathrm{c_i})$ of each individual i can be represented as:

$$s_i = s_{\Pi} * \frac{\Pi_i}{Y_i} + s_W * \frac{W_i}{Y_i}, \textit{and} \tag{12}$$

$$c_i = c_{\Pi} * \frac{\Pi_i}{Y_i} + c_W * \frac{W_i}{Y_i} \tag{13}$$

If the individual saving rate increases as individual relative income increases and women have a higher marginal propensity to consume than men – this hypothesis is consistent with evidence from Badru (2018), an inverse relationship may exist between gender inequality and aggregate consumption, such that, $\frac{\partial C}{\partial GG} > 0$. Also, women's labor force participation (LP) may impact their consumption expenditure either directly or indirectly such that $\frac{\partial C}{\partial LP} > 0$.[5]

Post-Keynesian models hypothesize that there is a dual effect of household debt on aggregate consumption (Dutt 2006; Palley 2010). This dual effect of household debt on aggregate consumption (e.g. see Duesenberry, 1949; Davanzati and Pacella, 2010; Ryoo and

[5] Women's labor force participation (LP) may impact on their consumption expenditure either directly (because they are able to earn their own income by working) or indirectly (because greater engagement of women in the labor market may imply that they are better able to organize in unions that advocate for higher wages for women, relative to men).

Kim, 2014; Setterfield, Kim, and Rees, 2015: Setterfield and Kim, 2016) can be represented by two terms in our consumption function. The first term represents the positive effect of increased borrowing on disposable income and therefore consumption. The second term represents the negative effect of the cost of servicing the debt on consumption. Denoting household debt as DH and the cost of servicing the debt as *INT* (interest rate).

The positive effect of increased borrowing on consumption is such that $\frac{\partial \mathrm{C}}{\partial \mathrm{DH}} > 0$, while the negative effect of the cost of servicing the debt on consumption can be represented as $\frac{\partial C}{\partial INT} < 0$. In such a setting, the overall effect of changes in households' indebtedness on consumption is ambiguous ($\frac{\partial C}{\partial DH} = \frac{\partial C}{\partial DH} + \frac{\partial C}{\partial INT}$) such that if the income effect is greater than the interest effect, i.e. $\frac{\partial C}{\partial DH}$ (income effect) > $\frac{\partial C}{\partial INT}$ (interest effect), the overall effect will be positive, and vice versa.

Our consumption function can thus be redefined as:

$$C = c * \left(\Pi + W_{m(t)} + W_{f(t)} + DH\right) - I(DH) + \frac{\partial C}{\partial GG} * GG + \frac{\partial C}{\partial LP} * LP \qquad \text{(2a)}$$

Equation (2a) extends the Keynesian aggregate consumption function in (2) where c is the average propensity to consume. Y is expected to have a positive effect on aggregate consumption. W is also expected to have a positive effect on consumption (e.g. Onaran and Galanis, 2012 and Hein and Vogel, 2007).

1.3. INVESTMENT

We start by defining the total investment (I) as the sum of business investment (I_b)and household investment (I_h):

$$I = I_b + I_h \qquad \text{(14)}$$

Business investment (I_b) is influenced by expected future demand, which can be proxied by the current level of output (Y). A common feature in the work of Keynes and Kalecki is the role attached to firms' rate of investment as a crucial determinant of output and employment.

These investment decisions by firms are assumed to be influenced by their expectations of future demand, given their existing stock and their ability to finance such investment decisions through such means as procuring external debt or utilizing internal cash flow. This ability to finance investment independent of saving is an important theoretical precondition for both Keynes and Kalecki.

Some of these features influencing the investment decisions of firms – together with households' investment decisions – are summarized in equation (3), above. Aggregate investment is then expected to be comprised of household and business investment. Y is expected to have a positive effect on I.[6] Similar to Stockhammer and Wildauer (2016), the effect of WS on I is ambiguous ($\frac{\partial I}{\partial WS} > 0 \ or \frac{\partial I}{\partial WS} < 0 \ or \frac{\partial I}{\partial WS} = 0$); while an increase in the wage share may indicate a loss of profits to capitalists or businesses – which may discourage future investment – such a change in WS may promote residential investment – where workers are home owners. It is expected that the effect of DH on investment in the long run is ambiguous because the accumulation of consumer debt results in a shift in the income distribution toward rentiers, who have a higher propensity to save (Asimakopulos, 1983). In a similar vein, an increase in DB is expected to have an initial positive effect on investment unless businesses incur further debt to service their outstanding financial obligation. The long-term interest rate (INT) is expected to have a negative effect on investment.

The net effect of higher female wages on profits and thus investment may be positive or negative, depending on the economic environment. The degree of firm mobility, for example may determine the impact of higher female wages on investment.[7]

NET EXPORTS

In the macroeconomic literature, changes in real foreign income (FY) and the nominal effective exchange rate (EX) are commonly assumed to be crucial determinants of a nation's export position. As in the case of investment, changes in the wage share (WS) represent redistribution from capitalists to workers, or vice versa and, as such, affect production costs and, accordingly, exports. In addition, wages affect the price level in an economy which further influences its international competitiveness. A priori, a negative association between

[6] An increase in Y represents an increase in demand and the level of production which, in turn, is likely to boost demand for capital and thus lead to greater investment.
[7] High firm mobility describes firms with low and limited sunk costs – including training costs – and easy firm entrance and exit. Mobile industries tend to be labor-intensive manufacturing firms as well as services, such as informatics, data processing, and possibly tourism (Seguino 1997, 2007).

WS and NX is hypothesized. We should note early on that to better represent our relationships between interest and the varying effect of gender and functional income distribution, separate import and export functions are estimated in this study. In the imports model, the potential dependence of export goods production on imported production inputs are accounted for by including exports as a right-hand-side variable in the import function.

The RW variable is used here to indicate not only the wage position of women relative to men but also their participation in the labor market—further represented by the inclusion of LP. While there is an obvious link between RW and NX using the same logic for the association between WS and NX, (LP), which is expected to lead to higher wages in the long run due to increased influence and bargaining power in the labor market, also has significant effects on the export sector, as is evident in the literature (Seguino 1997; Berik 2000; Blecker and Seguino 2002).

Ertürk and Darity (2000) highlight the dual effects of higher LP. Firstly, increased labor force participation by women often has the result of reducing both the time spent on unpaid caring labor—especially in the presence of rigid gender roles—and fertility rates, leading to increased earnings for women with potential negative impacts on the labor force, which, according to feminist theory can also be referred to as a produced means of production. Secondly, a lower wage position for women has a potentially positive impact on the composition and direction of production and thus exports, though which of these two cases has a higher effect on NX is expected to depend on the level of economic development [or structure] of a country.

As explained above, labor market conditions for women relative to men are expected to have different consequences in SIEs and LIAEs. For example, in LIAEs male labor is more concentrated in cash-crop production and nontradables (Collier, Edwards, and Roberts 1994; Seguino 2010a). Seguino and Were (2014) highlights that women, on the other hand, are more involved in subsistence agricultural production and sales; therefore, their involvement in this portion of production has no direct or immediate effect on international trade, as it does not directly influence global mobile investment, which is the opposite case for SIEs (as explained in the preceding subsection). An increase in women’s wages (holding men’s wages constant) will therefore be beneficial for domestic demand, at least in the short run.

Empirical research on intra-household bargaining suggests that men and women have different marginal propensities to consume, and these differences often depend on the type of consumption; for example, studies by Guyer (1988), Agarwal (1997) and Haddad (1999) find that men spend a larger proportion of their income on luxury goods rather than on basic household goods, while the opposite is true for women. This evidence appears to hold for countries at various levels of development. This could imply that for countries at a lower stage of development (e.g., LIAEs), consumption of luxury goods could be more import-intensive and where this is combined with a greater gender wage gap, household consumption on food, education and health care may be lacking, which may be detrimental for short-run domestic demand and long-run productivity growth.

The long-run version of this model looks more like a standard neoclassical supply-side model in that the only drivers of long-run growth are labor supply and productivity growth.[8] For both SIEs and LIAEs, labor supply growth is positively correlated with increases in female incomes, as increased female labor force participation is correlated with less gender inequality (Blau and Kahn, 2009; Seguino, 2012). However, Kaleckian models describe the long run as a succession of many short-run periods and, as such, short-run effects on long-run growth or the quasi-equilibrium are expected.

CUMULATIVE EFFECT ON INCOME

The effect of consumption, investment, and net exports on aggregate income can then be analyzed by substituting equations (2) – (4) into equation (1). Following Stockhammer, Onaran, and Ederer(2008) and, more closely, Stockhammer and Wildauer (2016), the impact of a change in WS on Y can then be represented as:

$$\frac{dY}{dWS} = \frac{g_1}{1 - g_2} \tag{15}$$

where $g_1 = \left(\frac{\partial C}{\partial WS} + \frac{\partial I}{\partial WS} + \frac{\partial NX}{\partial WS}\right)$ and $g_2 = \left(\frac{\partial C}{\partial Y} + \frac{\partial I}{\partial Y} + \frac{\partial NX}{\partial Y}\right)$

where g_1 represents the initial short-term effect and is defined as the resulting change in the level of aggregate demand due to a change in the functional income distribution level, $\frac{g_1}{1-g_2}$ represents the multiplier effect and g_2 represents the marginal effects of Y on its components. Therefore, $g_1 > 0$ implies an initial positive impact of a higher WS on private demand; in

[8]Here we refer to the inclusion of gender inequality indicators in a long-run productivity-driven growth model as is common in the mainstream literature, e.g., Dollar and Gatti (1999) and Klasen (2000). We, however, do not focus on this neoclassical model in this paper—especially the aspect of technology.

this case, the economy is wage-led (and vice versa). In the section below, we describe our method of estimating the gender effects.

GENDER EFFECTS

Research over time has shown that the persistent lower wages of women relative to men is a result of labour market discrimination and occupational segregation (Treiman and Hartman 1981; Padavic and Ross 1992; Reskin and Ross 1992). Also, gender-specific discrimination against women in the labour market often manifests differently in developed and developing nations. For example, we expect that gender inequality in developed countries can be best identified through the differences between the wages of men and women, while in developing nations, differential access to wage employment may be a better indicator of gender-based discrimination in the labour market (Collier et al. 1994; Elson, 1999; Cuberes and Teigner 2016). However, for this study, we are limited by data availability and, as such, specify the female-male wage ratio as the main gender inequality measure across our different country groups. The differentiation of gender effects across our different country groups is important as the effects of gender equality on macroeconomic variables are likely to depend on the structure of the economy (Seguino 2000).

We then account for the cumulative gender effects, in terms of calculated elasticities, on the macroeconomic aggregates. Specifically, we expect that an increase in gender equality (described here as gender wage equality) that results in injections exceeding leakages ($S + M < I + X$) is expansionary i.e., a redistribution stimulates aggregate demand, leading to an increase in output and employment (Seguino 2010a). A redistribution with this effect would be 'gender cooperative', which refers to a redistributive effect of increased female wages on firms and the ratio of female to male wages, and otherwise 'gender conflictive' when such an increase in the female-male wage ratio also results in a decline in male employment, thereby triggering an economic contraction (Seguino and Setterfield 2010).

Building on the Bhaduri and Margin (1990) model of demand-led growth employed in our analysis and extending it to incorporate gender wage inequality in developed economies, SIEs and LIAEs, we also compute the relative contribution of gender inequality in wages to actual growth by separating the relative growth effect of gender wage equality on the different components of aggregate demand.

For this analysis, the relative contributions of gender equality to economic growth can then be estimated by multiplying the estimated coefficient for each component, i.e. C, I and NX, with the actual change in RW (and LP), e.g. $\hat{\beta}_{CRW}\Delta RW$ for consumption (i.e. $\hat{\beta}_{CRW} = \partial C / \partial RW$), for each of the estimations in this study. This approach modifies the existing Kaleckian growth models by incorporating the effect of gendered labour market outcomes on the functioning of the macroeconomy in a similar manner as hitherto posited between class relations and the market economy. This may then imply that gender relations and the market economy, which are both independent and interacting domains, are part and parcel one of the other as the macroeconomy is itself gendered by way in which economic agents function within institutions.

III. DATA AND METHODS

In this section, the empirical relationship between personal and functional income distribution and aggregate demand is presented. Section 3.1 presents an overview of the data employed in the analysis; section 3.2 outlines the proposed empirical methodology.

1.4. Data

For this study, we compile an unbalanced panel dataset comprised of 46 countries using annual data over the period 1985-2011. The 46 countries, presented in the appendix, are selected based on data availability and are collected from all geographic regions. This dataset is particularly large in comparison to the existing literature on the empirical determination of demand-led growth regimes.

For further comparative analysis, we group our panel into high-income, middle-income and low-income countries, based on the World Bank classification, resulting in three panel datasets: 10 countries from 1985 to 2015 (Panel 1) for low-income, 12 countries from 1985 to 2015 (Panel 2) for middle-income and 24 countries from 1985 to 2015 (Panel 3) for high-income.

Annual data for this study were derived from the Bank of International Settlements (BIS), IMF Database, the World Development Indicators (WDI) of the World Bank, Federal Reserve Economic Data (FRED) and the International Labour Organization (ILO) data banks.

A summary of the definition and sources of the variables in the macro panel analysis is presented in Table 5.2 below.

Following the relatively long timespan (T = 31) of our dataset, we need to consider several potential econometric issues often associated with panel data of this magnitude, namely, cross-sectional dependence (CSD), dynamics, slope heterogeneity and the presence of unit roots.

1.5. Methodology and Preliminary Tests

To examine the nexus between income distribution and aggregate demand, a framework based on Bhaduri and Marglin's (1990) formulation of the Post-Keynesian demand-led growth model is proposed (Keynes 1936; Kalecki 1939, 1971). The adopted theoretical model for this analysis follows the approach by Stockhammer, Onaran, and Ederer (2008), and adds to this the gendered impact of income inequality within a panel framework.

We therefore provide a formal statistical analysis of distributional effects on aggregate demand following standard Keynesian representations as observed in equations (1) through (5), above. To shed light on possible heterogeneity of the effects across countries at various stages of development, reporting separate results for high-income, middle-income and low-income countries or, alternatively, advanced economies, SIEs, and LIAEs. This is done to account for the role that the economic structure plays in influencing the relationship between gender equality and macroeconomic outcomes, and the nature of persistence of this association.

Section 2.2 above outlines the theoretical model for the empirical estimation, particularly, the determination of the multiplier effect and the implication of the numerator (private excess demand) and denominator terms (see equation [5]). To infer g_1, we calculate the sum of the marginal effects of WS on C, I, and NX. To proceed with our empirical analysis, one needs to consider the issues of non-stationarity, dynamics, heterogeneity, and CSD that may emerge from a macro panel dataset of this magnitude.

IV. Stationarity

Before testing for long-run cointegration between our variables of interest, we first check for the order of integration in our series using unit root tests. Specifically, we conduct panel data unit root tests allowing for homogenous and heterogeneous slopes. We employ the Levin-Lin-Chu (LLC) unit root tests by Levin, Lin, and Chu (2002), which impose a homogeneity assumption, and the first-generation Im-Pesaran-Shin (IPS) test by Im, Pesaran, and Shin (2003) to allow for heterogeneous autoregressive coefficients; both tests assume a standardized average of individual Augmented Dickey-Fuller (ADF) statistics to test the pooled null hypothesis of a unit root against a heterogeneous alternative. We also employ a second-generation unit root test (the cross-sectionally augmented IPS [CIPS] of Pesaran [2007])) to additionally account for CSD among panel countries.

To test for a unit root, consider the conventional univariate ADF specification as follows:

$$\Delta y_{it} = \alpha_i + \beta_i t + (\rho_i - 1) y_{it-1} + \sum_{j=1}^{\rho} \delta_{ij} \Delta y_{it-j} + u_{it} \tag{16}$$

with $H_{0:}\, \rho_i - 1 = 0, i = 1, \dots, N$

$$H_{1:}\, \rho_i - 1 < 0, i = 1, \dots, N_1;\ \rho_i - 1 = 0, i = N_1 + 1, \dots, N$$

where y_{it} represents the respective panel series considered for country *i* at time *t*. Under the LLC test, the null hypothesis is that $\rho_1 = \rho_2 = \cdots = \rho_N = 1$ while the alternative hypothesis assumes that $\rho_1 = \rho_2 = \cdots = \rho_N < 1$.

Where our panel series are I(0) or I(1), standard autoregressive distributed lag (ARDL) panel regressions can be relied on to produce efficient estimates, as demonstrated in a series of papers by Pesaran and others (Pesaran and Smith 1995, Pesaran 1997, and Pesaran, Shin and Smith 1999).

Appendix 2 presents the result of the LLC, IPS and CIPS panel unit root tests. The various tests produce conflicting results on the order of integration. Using the LLC unit root test, all variables, other than imports, exchange rates, interest rates and wage share, possess a unit root. Results from the IPS test, which allows for heterogeneous coefficients across countries, shows that all variables other than the female-to-male wage ratio, interest rates and exchange rates to be non-stationary in levels. However, these variables become stationary upon first differencing. The IPS test uses the simulated critical values provided by Im et al (2003) for different *N* cross-sectional units and *t* time-series.[9]

[9] We use EViews for the unit root tests which employs the relevant critical values, or linearly interpolated values, in evaluating the significance of the test statistics.

In addition to the LLC and IPS tests, we employ the CIPS test, using the simulated critical values of CIPS listed in Pesaran (2007), to allow for correlation between the error terms across the panel of countries. For some of the variables, we find dissimilar results from the LLC and IPS tests at level; given that the CIPS test takes account of cross-sectional dependence – which is likely to be an issue among some of our variables following the recognition of cross-section dependence using the CD test in the latter sections of this chapter – it is possible that our results based on the IPS test are spurious. Upon differencing the relevant series however, none of the variables were found to be > I(1).

Based on the LLC, IPS and CIPS tests, some variables appear to be I(1) for all countries while others can be assumed to follow I(0) processes for some of the countries in the panel. Our CIPS test suggests that a larger number of the variables are I(0) series for at least some countries; as such, the LLC and IPS test results may not be entirely reliable due to the presence of cross-sectional dependence as observed in the next section.[10]

Cross-sectional Dependence and Slope Heterogeneity

Standard empirical estimators for panel data, such as the fixed effects and system GMM, commonly impose assumptions of independence across cross sections and slope homogeneity across countries. Recent research has however shown that these assumptions of zero covariance of the error terms and common slope parameters in panel estimations are easily contravened (Westerlund and Edgerton 2008; Eberhardt and Teal 2011; Sarafidis and Wansbeek 2011).

Furthermore, given our panel, which consists of countries at different stages of development with differing social and economic conditions and, as such, conceivable structural differences, panel data methods that rely on pooling the data may result in potentially inconsistent and misleading estimations (Pesaran, Shine and Smith 1999). It is therefore of importance that, in the presence of cross-sectional error dependencies and cross-country heterogeneity, an estimation strategy that adequately accounts for these features is employed. Disregarding such dependencies could lead to spurious inference and substantial bias in the estimated parameters of the specified model. It is also necessary to properly model dynamics to ensure that the estimated long-run effects are consistent.

[10] We base our final conclusions on the CIPS test results due to the issues stated above.

To this end, we carry out CSD tests under the null hypothesis of cross-sectional independence among countries, and poolability tests under the null of homogeneity of slopes across countries. To test for the existence of CSD among our countries, the general Pesaran (2004) cross-sectional dependence test (CD_p) is employed, which under H_0 is distributed as a standard normal distribution ($〚CD〛\_p \sim N(0,1)$ for $T\_ij > 3$ and sufficiently large N, as is the case in this study.

This test is robust to non-stationarity (as any spuriousness from stationarity issues will be observed from averaging the factors), parameter heterogeneity or structural breaks and is assumed to perform well in small samples (Pesaran, 2004).

The CD_p statistic can be defined as:

$$CD_p = \sqrt{\left(2T \Big/ \left(N(N-1)\right)\right)} \left(\sum_{i=1}^{N-1} \sum_{j=i+1}^{N} \hat{\rho}_{ij}\right) \qquad \textbf{(17)}$$

where $\hat{\rho}_{ij}$ is defined as the average pairwise correlation of the respective panel series with a null hypothesis of no CSD.

To test for slope homogeneity across countries, Pesaran and Yamagata's (2008) modified version of the Swamy (1970) poolability test ($\vec{\Delta}$ test) is employed and extended to the case of large N relative to T (as in this case) under homogeneity null under the condition of $(N,T) \to \infty$ without any restrictions on $\sqrt{N/T}$ with normally distributed error terms.

Following Pesaran and Yamagata's (2008) $\vec{\Delta}$ test for slope homogeneity in large panels, we first compute the standardized version of the Swamy (1970) poolability test given by:

$$\tilde{S} = \sum_{i=1}^{N} \left(\breve{\beta}_i - \acute{\beta}_{WFE}\right)' \frac{x_i' M_\tau x_i}{\acute{\sigma}_i^2} \left(\breve{\beta}_i - \acute{\beta}_{WFE}\right) \qquad \textbf{(18)}$$

where $\breve{\beta}_i$ is the estimator from the pooled OLS, and $\acute{\beta}_{WFE}$ is the estimator from the weighted fixed effects pooled estimations of the regression models derived from $(\mathbf{1} \sim \mathbf{4})$ above; x_i is a $k \times 1$ vector of regressors; M_τ is an identity matrix; and $\acute{\sigma}_i^2$ is the mean square error from the ordinary least squares (OLS) regression estimated for each cross-sectional unit.[11]

The standardised dispersion statistic $(\vec{\Delta})$ is then defined as:

[11] Refer to Pesaran and Yamagata (2008) for a detailed presentation of the modified Swamy (1970) test and the broader definition of the estimators given in equation (10).

$$\vec{\Delta} = \sqrt{N}\left(\frac{N^{-1}\tilde{S} - k}{\sqrt{2k}}\right) \tag{19}$$

Under the null of homogeneity, for all *i* under the condition $(N, T) \to \infty$, when $\sqrt{N/T} \to \infty$ and with normally distributed error terms, the $\vec{\Delta}$ test has an asymptotic standard normal distribution.

Table 1 reports the results from the Pesaran (2004) cross-sectional dependence test (CD_p). The Pesaran and Yamagata (2008) standardised version of the Swamy (1970) slope homogeneity tests $(\vec{\Delta})$ results are reported in Section 3. The cross-section dependence tests strongly indicate that we reject the null hypothesis of independent cross-sections is rejected in favour of the alternative of dependent cross-sections for the analysed panel series. Following the results for CSD, it is important that the presence of cross-sectional dependence be allowed for in our empirical analysis.

[Table 1 about here].

V. Empirical Approach

Our empirical model is specified within a heterogeneous dynamic panel data setting based on the Pesaran and Smith (1995) standard panel ARDL model (with p and q lags) with a multifactor error structure of the form:

$$y_{it} = \alpha_i + \sum_{j=0}^{q} \delta_{ij} x_{i,t-j} + \sum_{j=1}^{p} \lambda_{ij} y_{i,t-j} + u_{it} \tag{20}$$

$$\text{where} \quad u_{it} = \gamma_i' \boldsymbol{f}_t + \varepsilon_{it},$$

$$x_{it} = a_i + b_i y_{i,t-1} + \Gamma_i' f_t + v_{it}$$

where i = 1,2, …, N; t = 1,2, …, T. $x_{i,t}$ is a k × 1 vector of regressors for cross-section unit i at time t; δ_{ij} represents the k × 1 coefficient vectors; λ_{ij} are scalars; and α_i and a_i represents a set of country-specific fixed effects capturing the impact of unobserved country-specific time-variant heterogeneity.

ε_{it} and ν_{it} represent the idiosyncratic errors that are assumed to be independently distributed across i and t; with zero mean and constant variance; $\boldsymbol{f_t}$ is an m × 1 vector of unobserved common factors that capture cross-sectional dependencies across countries; and γ_i' and Γ_i' are country-specific 1 × m matrices of corresponding factor loadings.

This ARDL approach allows both the short-run dynamics, and the long-run [quasi-] equilibrium effects of income distribution and gender equality on each of the variables in the macroeconomic and net export equilibrium conditions to be separately identified. Furthermore, the ARDL model specified in equation (11) takes into account certain properties of the data series, such as non-stationarity and unobservable common factors.

Following Pesaran et al. (1998) and Pesaran et al. (2001), equation (33) can be expressed in an equivalent error correction form such that short-run dynamics are influenced by deviations from the equilibrium as denoted below:

$$\Delta y_{it} = \beta_i\left(y_{i,t-1} - \theta_i' x_{i,t-1}\right) + \sum_{j=0}^{q-1} \delta_{ij}'^{*} \Delta x_{i,t-j} + \sum_{j=1}^{p-1} \lambda_{ij}^{*} \Delta y_{i,t-j} + u_{it} \qquad \textbf{(21)}$$

Such that, $\beta_i = -\left(1 - \sum_{j=1}^{p} \lambda_{ij}\right)$; $\theta_i = {\sum_{j=0}^{q} \delta_{ij}} \Big/ {1 - \sum_{j=1}^{p} \lambda_{ij}}$; $\delta_{ij}'^{*} = -\sum_{m=j+1}^{q} \beta_{i,m}$;

$$\lambda_{it}^{*} = -\sum_{m=j+1}^{p} \lambda_{i,m}$$

where β_i is the error correction coefficient that captures the speed of adjustment of deviations to long-run equilibrium, θ_i captures the long-run equilibrium relationship between our dependent and explanatory variables, and $\delta_{ij}'^{*}$ and λ_{it}^{*} define the short-run dynamics. A cointegrating relationship is inferred from a negative and significant β_i, which also provides an indication of the stability of the long-run cointegrating relationship.

This ECM specification is preferred to static or more restricted dynamic models because it accounts for heterogeneity across countries by allowing differing β_i, θ_i and δ_i across the panel sample. It is also important to note that estimates of β_i are consistent whether the panel variables are I(0) or I(1) (Pesaran and Shin, 1999). Also, given that these models are autoregressive, they do not suffer from endogeneity bias if sufficient lags are included (Chudik and Pesaran, 2013).

Test results described in the previous section provide evidence of non-stationarity and CSD in our panel series. Therefore, the estimation of equation (12) for the theoretical models (Section 2) requires macro panel-data techniques that allow for non-stationary series and

CSD across countries while also providing estimates for all panel members. Another issue is that of endogeneity that arises from feedback between income distribution and aggregate demand in the specification of our theoretical models. In a seminal paper, Hansen and Tarp (2001) note that issues of endogeneity can be addressed to a reasonable measure with the use of dynamic panel estimation. By incorporating dynamics into equation (12), in this case as a lagged dependent variable, exogeneity restrictions are relaxed. This technique allows for weakly exogenous regressors within a dynamic panel model, as in equation (13). Additionally, it allows for the possibility of feedback between variables in the respective consumption, investment and net export equations; for example, past levels of the dependent variables can affect the regressors resulting in a case of weak exogeneity where such reverse causality exists.

Previously, we shed some light on the structural differences that may factor into the relationship between gender and the aggregate demand components, such as those between high-income/developed economies and low-income agricultural societies. These potential differences in gender behavior and outcomes from one country to another (depending on their level of development) may mean that the parameters of equations (1) – (4) are not equal for all panel countries. In such a case, panel regression techniques that require pooling of the panel series, such as the dynamic fixed effects (DFE) and generalized method of moments (GMM), may produce inconsistent, and possibly misleading, estimates. Following similar considerations, Pesaran and Smith (1995) develop a mean group (MG) approach, which allows for the separate estimation of the parameters for each cross-section; the mean of the estimated coefficients is then calculated to produce the reported coefficient estimates. Pesaran (2006) argues that error processes may suffer from cross-sectional dependence when panel cross sections are affected by factors not included in the estimation process or when there are spatial spillovers. Where such common factors are not adequately controlled for, misleading estimates may result.

Motivated by these concerns, we employ the Cross-Sectional Autoregressive Distributed Lag (CS-ARDL) model of Chudik and Pesaran (2013). As an additional robustness check, we use a Cross Sectional – Distributed Lag (CS-DL) model (a reformulated autoregressive distributed lag ARDL specification) to help avoid possible bias in the long-run estimates resulting from bias in the parameter of the lagged dependent variable. Both approaches allow for country-specific heterogeneity, error variances, and cross-country correlations. They are robust to endogeneity created by unobserved common factors, as well as omitted variable bias. The main advantage of the CS-DL regression is that it yields more precise long-run estimates than CS-ARDL when the time dimension of the data is not sufficiently long (T<50), as is the case in this study.

Common Correlated Effects Mean Group Estimator

We estimate our goods market behavioral equations (equation (1) ~ (4)) building on insights from multi-factor models in nonstationary panels (Kapetanios et al., 2011; Pesaran, 2006). To estimate our relevant average long-run effect and short-run dynamics, we employ the CS-ARDL approach and CS-DL approach proposed by Chudik et al. (2016).

For the CS-ARDL model, Chudik and Pesaran (2013) develop the Common Correlated Effects (CCE) method of Pesaran (2006) by showing that an ARDL model can be augmented with cross-section averages of the observable variables to account for the unobserved common factors (f_t). equation (13) to (16) are estimated following Chudik and Pesaran (2013) by augmenting the ARDL model (equation (11)) with cross-sectional averages of the model's observable variables.

Assuming the regressors are independently distributed of the slope estimates, we can (as in Chudik and Pesaran (2013) substitute for u_{it} while averaging Eqs. (11) and (12) across i, we have:

$$f_t = \bar{\phi}^{-1}\left(\bar{y_t} - \bar{\alpha} - \sum_{j=0}^{q} \bar{\delta_j}\, \overline{x_{t-j}} - \sum_{j=1}^{p} \bar{\lambda_j}\, \overline{y_{t-j}} + \bar{\varepsilon}_t\right) \quad \textbf{(22)}$$

where $\bar{\phi} = N^{-1}\sum_{i=1}^{N}\phi_i$; $\bar{x}_{t-j} = N^{-1}\sum_{i=1}^{N}x_{it}$; $\bar{y_t} = N^{-1}\sum_{i=1}^{N}y_{it}$; $\bar{\lambda}_j = N^{-1}\sum_{i=1}^{N}\lambda_i$; $\bar{\alpha} = N^{-1}\sum_{i=1}^{N}\alpha_i$; $\bar{\delta_j} = N^{-1}\sum_{i=1}^{N}\delta_i$, $j = 0, 1, \ldots, q$;

For $N \rightarrow \infty$ and $\phi \neq 0, \bar{\varepsilon}_t = 0,$ and cross-sectional correlation can be controlled for via a linear combination of the cross-sectional averages of the dependent and independent variables.

Modifying the model in Eqs. (11) and (12) accordingly, we obtain:

$$\bar{y_t} = \bar{\alpha} + \sum_{j=0}^{q} \bar{\delta_j}\, \overline{x_{t-j}} + \sum_{j=1}^{p} \bar{\lambda_j}\, \overline{y_{t-j}} + \bar{\varepsilon}_t \quad \textbf{(23)}$$

where $\bar{\varepsilon}_t = N^{-1} \sum_{i=1}^{N} \varepsilon_{it}$

The corresponding CS-ARDL specification of equation (11) is then given by:

$$y_{it} = \alpha_i + \sum_{j=1}^{p} \lambda_{ij} y_{i,t-j} + \sum_{j=0}^{q} \delta_{ij} x_{i,t-j} + \sum_{j=1}^{p} \psi_j \overline{y_{t-j}} + \sum_{j=0}^{q} \kappa_i \overline{x_{t-j}} + \varepsilon_{it} \quad \textbf{(24)}$$

Following Chudik and Pesaran (2013), we then estimate equation (15) using the MG estimator of Pesaran and Smith (1995). The corresponding CS-ARDL ECM specification of equation (15) can be presented thus:

$$\Delta y_{it} = \alpha_i + \beta_i \left(y_{i,t-1} - \theta_i' x_{it} \right) + \sum_{j=0}^{q-1} \delta_{ij}^{*} \Delta x_{i,t-j} + \sum_{j=1}^{p-1} \lambda_{it}^{*} \Delta y_{i,t-j} + \sum_{j=1}^{p} \psi_j \overline{y_{t-j}} \quad \textbf{(25)}$$

$$+ \sum_{j=0}^{q} \kappa_i \overline{x_{t-j}} + \sum_{j=1}^{p-1} \Phi_j \overline{\Delta y_{t-j}} + \sum_{j=0}^{q-1} \Gamma_i \overline{\Delta x_{t-j}} + \varepsilon_{it}$$

We can then proceed to the Mean Group (MG) estimation of equation (16) which provides consistent estimates of the model parameters; this approach is also robust to weak exogeneity. Also, the choice of a dynamic model has the additional advantage of providing both long-run estimates and short-run dynamics.

Cross-Sectionally Distributed Lag Approach

As earlier stated, we employ the traditional ARDL and CS-DL methods as sensitivity checks. The CS-DL method recently proposed by Chudik et al. (2016) directly estimates the long-run coefficients in a dynamic panel model. In this case, no short-run dynamics are provided from the MG estimation in the distributed lag representation of the model and only a truncation lag (ρ) is needed for this estimation.

Controlling for common factors and cross-sectional dependence, we augment our initial ARDL model with cross-sectional averages of dependent and independent variables rewritten in a distributed lag presentation of the model. This CS-DL specification by Chudik et al. (2016) can then be rewritten as:

$$y_{it} = \alpha_i + \theta_i' x_{it} + \sum_{j=0}^{q} \delta_{ij} \Delta x_{i,t-j} + \sum_{j=0}^{p_{\bar{y}}} \psi_{ij} \overline{y_{t-j}} + \sum_{j=0}^{q_{\bar{x}}} \kappa_{ij} \overline{x_{t-j}} + \varepsilon_{it} \quad \textbf{(26)}$$

There are several advantages to employing each of the three different approaches and some of these have been highlighted. In truth, we favor the CS-ARDL approach particularly for its autoregressive properties and its ability to tackle some of the problems related to cross-sectional dependence and cross-country heterogeneity while also allowing for lagged regressors of the dependent variable; this is the main reason for our choice of the CS-ARDL over the CS-DL approach.

This clarification is important because it can be argued that the CS-DL method is most reliable in terms of issues regarding sampling uncertainty; following results from Monte Carlo simulations, Chudik et al. (2013; 2016) posit that the ARDL and CS-ARDL techniques are more susceptible to large sampling errors with smaller samples. By ignoring the short-run dynamics during estimation, the CS-DL technique is also able to by-pass some of the issues of estimation performance that can result from lag length misspecification (Chen and Vujic, 2016). However, we are interested in the autoregressive term and the short-run dynamics in this chapter, so the CS-ARDL is more suited to our empirical purpose.

VI. EMPIRICAL RESULTS

Using pooled time-series cross-section data (with N = 45 and T = 31, comprising 1097 observations) on aggregate income (Y), aggregate private consumption (C), aggregate private investment (I), net exports (NX), adjusted wage share (WS), female-to-male wage ratio (RW), long term interest rate (INT) and the effective exchange rate (EX), we estimate consumption, investment and net exports equations in a dynamic heterogeneous panel.

We begin our analysis by choosing the optimal lag structure (p,q) for the ARDL, CS-ARDL and CS-DL models. Chudik et al. (2016) explain the need for an appropriate lag length, to ensure that the ARDL estimates are consistent, while also stating that where longer lags than necessary are employed, estimates with poor sample properties may result. We apply the Schwartz Bayesian Criterion (SBC) to select the lag length.[12] Specifically, we employ two approaches to model selection using SBC. Following Loayza and Ranciere (2006) and Kim et al., (2016), we select the optimal lag structure for each model on a country-by-country basis, from which the most common lag length among the countries is selected as the most

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[12] We also use the Akaike information criterion (AIC) as a cross-validation for the SBC method; we find that the most common lag length on the country-specific lag order is the same for both the AIC and SBC except in the case of the investment function. Eviews software is used in this model selection process.

appropriate for the model estimation. As an additional sensitivity check for model selection, we employ a VECM estimation on the entire panel and the optimal lag is selected based on the lowest values of AIC and SBC. In both cases, we impose a maximum lag order of three in our specifications and find a consistent lag choice using the Schwartz Bayesian Criterion.

Our results show that the ARDL (1, 1, 1) and CS-ARDL (1, 1, 1) models achieves the lowest SIC value and is thereby chosen as the best model. We then proceed to conduct our empirical estimations using the MG approach for all models. As proposed in Section 5.1.3 above, we estimate the error-correction model for all ARDL and DL specifications by imposing the same lag structure selected by the criterion discussed above. We also consider p=2 to check for sensitivity of our results to lag length specification.

Having already conducted formal tests to examine the properties of CSD and non-stationarity (unit roots) for our panel data set and confirmed the presence of these issues, we proceed to estimate the heterogeneous dynamic ECM using the ARDL, CS-ARDL and CS-DL estimators, results of which are reported in Appendix 1 and 2. It should be noted that while we report and discuss results for the ARDL, CS-ARDL and CS-DL specifications, we base our analyses on results from the CS-ARDL model due to the possibility of endogeneity with some of our explanatory variables, and because of the presence of CSD. Moreover, Post-Keynesian macroeconomists consistently call attention to the role of the class distribution of income in influencing short-run outcomes which have implications for the longer run. This feature of the discourse makes the CS-ARDL estimation especially useful in this study as it provides the long-run estimates and short-run dynamics of the system.

Furthermore, the favorable results and relevant diagnostics (Root Mean Square Error (RMSE), *CSD* test statistic) obtained enable us to conclude the latter part of the demand effect estimations using results based on the CS-ARDL models augmented with three lags of cross-section averages.[13] In addition, due to the small sample bias the ARDL approaches face, the ECM is computed using a jack-knifed bias correction procedure on our estimates. This Jack-knife procedure (originally proposed by Quenouille (1949) and further developed by Tukey (1956)) serves to correct for bias by a method of [re]sampling without replacement. Specifically, for this study, the “half panel” jack-knife by Chudik et al., (2016) expressed below is employed:

[13] Chudik and Pesaran (2015) advise that the inclusion of sufficient cross-sectional lags (preferably 3 cross-sectional lags) is necessary to ensure allowance for the possibility of cross-sectional error correlations due to omitted common effects.

$$\hat{\pi}^{J}_{MG} = 2\hat{\pi}_{MG} - \frac{1}{2}\left(\hat{\pi}^{a}_{MG} + \hat{\pi}^{b}_{MG}\right) \quad \textbf{(27)}$$

where $\hat{\pi}^{a}_{MG}$ is the mean group estimate of the first half (t =1, …, $\frac{T}{2}$) of the panel and $\hat{\pi}^{b}_{MG}$ of the second half (t = $\frac{T}{2} + 1$, …, T) of the panel. This time-series bias correction is carried out in Stata.

Moreover, we assume nonzero effects for our outcome variables. Evidence suggests that when is negative and statistically significant and a long-run cointegrating relationship is established under a common correlated effects cointegration technique, it is impossible to distinguish a null effect from a very small effect (Muller 2004; Lane et al. 2015 Choi and Chudik 2019). Following this evidence, negligible effects, even when not statistically significant, are considered in estimating the cumulative distributional effects and in proposing gender and distributive effects.

For all the aggregate demand components, the baseline model is subjected to two robustness checks. The first check entails using the ARDL and CS-DL approaches as sensitivity checks for our CS-ARDL estimates. As a second check we expand the baseline specifications in each function by adding relevant explanatory variables using the ARDL, CS-ARDL and CS-DL. Furthermore, for all specifications, we only report the first lag short-run coefficients (e.g. $\partial\Delta Y_{it}/\partial\Delta X_{it}$). Our short-run results are only at times statistically significant, and especially so with the first lag results. Coefficients from higher order lags are rarely ever significant and when they are they very small magnitudes.

1.6. Consumption

Following the literature on demand-led growth, the baseline consumption function is estimated where C_t is a function of the income variable, Y_t , the variable capturing gender wage inequality, RW_t and the wage share variable, WS_t - with household debt-to-GDP (DH_t) and female labor force participation (LP_t) included as an additional control – based on a panel model of the general form:[14]

[14]Assuming a linear relationship is a good starting point for our analysis as most of our key explanatory variables (e.g., WS and RW) vary over only a narrow range of values. We then continue our analysis with a log-log specification as we expect that the impact of the independent variables on the dependent variables may in some cases be different from one level to the other. Furthermore, previous studies (Stockhammer, Onaran, and Ederer2008; Onaran and Galanis 2012) also employ a log-log functional form given the interest in calculating

$$lnC_{it} = f(\, lnY_{it}, lnRW_{it}, lnWS_{it}, lnDH_{it}, lnLP_t) \tag{28}$$

The ARDL, CS-ARDL and CS-DL estimates of equation (19) are reported in Table 2 with jack-knifed standard errors.[15] As earlier stated, the CS-ARDL estimator is preferred over others for this study due to compelling evidence of cross-sectional dependence and the possible endogeneity issues with our variables.

VII. Estimates of Long-run effects

The least squares estimates obtained from the panel ARDL and DL specifications in Table 5.5 report the results for the baseline specification and Table 5.6 shows the results for the extended regression when an additional explanatory variable (DH) is included. Each panel gives the error correction variant of the Mean Group (MG) estimates of the long-run effects on aggregate consumption from changes in our explanatory variables.

Pesaran and Smith (1995) posit that the MG estimates are consistent under fairly general conditions where the errors are cross-sectionally independent. We do however observe that CSD is present under the ARDL specifications, but not with the CS-ARDL model. We also observe that the error correction coefficient is negative and statistically significant in all the ARDL specifications (with the exception of the CS-ARDL (2,2) model), indicating a mean reversion to a non-spurious long-run relationship, thus implying cointegration between our variables.

Table 2: Regression results for the Consumption function

	ARDL		**CS-ARDL**		**CS-DL**
Lags	(1,1)	(2,2)	(1,1)	(2,2)	ρ =1
Regressions with Key variables					
$\hat{\boldsymbol{\theta}}_{Y}$	0.991***	0.938***	0.809***	0.633**	0.529***
	(0.080)	(0.009)	(0.177)	(0.220)	(0.059)
$\hat{\boldsymbol{\theta}}_{RW}$	0.111**	0.016	0.992	-0.856	0.048
	(0.166)	(0.031)	(0.800)	(2.044)	(0.342)
$\hat{\boldsymbol{\theta}}_{WS}$	0.331**	0.210***	0.494**	0.831*	0.035

marginal effects for the C, I, and NX models—which requires estimated elasticities. Finally, this transformation allows for easier interpretation of our estimates.

[15] All estimations for the AD components equations are carried out using the Stata 13 software.

	(0.169)	(0.031)	(0.036)	(0.223)	(0.265)
$\hat{\lambda}$	-0.413***	-0.174***	-0.549***	0.014	n.a.
CD	84.00***	92.49***	-0.50	-0.24	-0.69
$\vec{\Delta}$	16.857**				

Notes: $\hat{\theta}$ is an indicator of the long-run coefficient on regressor $x_{i,t}$ in equation (33).

$\hat{\lambda}$ is the speed of adjustment (ECT).

The CD test reports the CD statistics, instead of p-values.

ρ is the truncation lag.

$\vec{\Delta}$ reports the Pesaran and Yamagata (2008) poolability test results for the estimated function.

Jack-knifed standard errors are reported in parentheses.

n.a. stands for not applicable when there are not enough observations to conduct the relevant estimations when using certain lags to deal with the potential serial correlation and CSD issues.

***, ** and * denote significance at the 1%, 5% and 10% levels, respectively.

With regards to the baseline specification, table 2 reports closely similar ARDL and CS-ARDL estimates for the long-run income effect on consumption, with highly significant elasticities varying from 0.63 to 0.99. These income elasticity results for the ARDL and CS-ARDL estimates appear more comparable under identical lag structures; when ρ=1 (i.e., ARDL [1,1] AND CS-ARDL [1,1]), the income elasticity of consumption ranges from 0.81 to 0.99. The long-run average coefficients on the wage share are positive under all model specifications, consistent with the demand-led growth literature (Stockhammer, Onaran, and Ederer 2008; Onaran and Galanis 2012; Hein and Vogel 2007. In addition, we observe that the wage share coefficients are statistically significant under all but the DL (1,1,1) model. Using our preferred CS-ARDL results, we expect that in the long run, a 1 percent increase in the wage share is associated with a 0.49 percent increase in aggregate consumption.

The coefficient of the female-to-male wage ratio is positive under all specifications (except for the CS-ARDL estimates when ρ = 2). It is worth noting that the RW coefficients are positive and significant only in the ARDL specification. The CS-ARDL and DL estimates, while mostly positive, are not significant and as such do not provide robust evidence to support the long-term relationship between aggregate consumption and gender wage equality. In this case, our ARDL (1,1,1) estimate suggests a 0.11% response of aggregate consumption to a 1% improvement in the proportion of the female wage bill.

1.1.1.1.1. Robustness to additional explanatory variables

start from here

Our baseline specification above imposes some simplifying assumptions on the consumption function. We therefore attempt to relax these assumptions by expanding the model estimated in Table 2 above. Specifically, we consider the household debt-to-GDP ratio (DH) and the female labour force participation (LP) as additional factors that can potentially influence the aggregate level of consumption. We are constrained from adding more controls because of reduced degrees of freedom.

We find the income elasticity of consumption estimates to fall within a range of 0.72 to 1.03 when $\rho = 1$; all estimates are significant as in the baseline model, following *a priori* expectations. The ARDL, CS-ARDL and DL long-run coefficients on the wage share are all positive. However, unlike the baseline results, only the CS-ARDL (1,1) estimates are statistically significant with a coefficient of 1.168.

Table 5.1: Regression results for the Consumption function (equation (23))

	ARDL		**CS-ARDL**		**CS-DL**
Lags	(1,1)	(2,2)	(1,1)	(2,2)	ρ=1
Regressions with additional variables					
$\hat{\boldsymbol{\theta}}_Y$	0.994***	0.908***	1.027***	1.095***	0.717***
	(0.161)	(0.116)	(0.144)	(0.197)	(0.108)
$\hat{\boldsymbol{\theta}}_{RW}$	0.792	0.362	0. 377**	-3.852	0.664
	(1.022)	(0.535)	(0.152)	(3.499)	(0.703)
$\hat{\boldsymbol{\theta}}_{WS}$	0.329	0.085	1.168*	4.348	0.121

	(0.403)	(0.643)	(0.674)	(3.803)	(0.115)
$\widehat{\boldsymbol{\theta}}_{LP}$	0.089	-0.069	0.178**	-0.083	0.048
	(1.063)	(0.765)	(0.568)	(0.137)	(0.046)
$\widehat{\boldsymbol{\theta}}_{DH}$	0.181	-0.141	0.006	na	0.097
	(0.128)	(0.011)	(0.055)		(0.068)
$\hat{\boldsymbol{\lambda}}$	-0.470***	-0.613**	-1.239	-0.747***	
CD	114.07**	86.23**	42.42***	42.41***	0.42*
$\vec{\Delta}$	11.347**				
Refer to Table 5.5 notes					

Turning to the long-run effects of RW, we find that all but the CS-ARDL (2,2) estimates suggest a positive relationship between the female-male wage ratio and aggregate consumption for our entire panel. However, only the one-lag CS-ARDL specification returns a positive and significant estimate for RW under the extended regression. More interestingly, we notice that when two lags are imposed on the CS-ARDL approach, a negative coefficient is observed, even though the estimates are not significant. While we cannot draw firm conclusions on this phenomenon (given that the estimates are not statistically significant in this case), it is however not unexpected that the persisting effect of a closing of the gender wage gap may negatively affect consumption, due to an increasing effect on saving after an initial period (Seguino and Floro, 2003).

Table 5.2: Short-Run Regression Results for the Consumption Function (equation (23))

	ARDL		**CS-ARDL**	
Lags	(1,1)	(2,2)	(1,1)	(2,2)
Regressions with Key variables				
$\widehat{\boldsymbol{\partial}'^*}_{Y}$	0.148***	0.324***	0.039*	0.315***
	(0.042)	(0.054)	(0.1.06)	(0.067)
$\widehat{\boldsymbol{\partial}'^*}_{RW}$	0.454	0.134	0.348	0.196
	(0.385)	(0.305)	(0.278)	(0.158)
$\widehat{\boldsymbol{\partial}'^*}_{WS}$	0.113	0.077	0.098*	-0.147**

	(0.204)	(0.229)	(0.207)	(0.059)

where $\widehat{\partial'^{*}}$ represents the first-lag short-run coefficients related to the respective regressors.

Furthermore, The ARDL, CS-ARDL and DL estimations suggest a positive relationship between female labor force participation rate (LP) and aggregate consumption when one lag is imposed and a negative association when ρ = 2. This finding (while only statistically significant in the CS-ARDL (1,1) specification) may imply little long-term effects of increased participation of women in the labor force on consumption.

For the additional DH variable, we do not find significant results under any of the specifications. These observed results are also not robust in comparison to the various specifications in terms of the direction of effect.

Short-run Dynamics

Table 5.7 and 5.8 report the short-run results of the MG estimations for the baseline and extended regressions respectively. The baseline CS-ARDL (1,1) short-run dynamics indicate that income and the wage share have a positive and significant effect on aggregate consumption. As expected, the RW coefficient is positive but not significant in the CS-ARDL case.

Table 5.3: Short-Run Regression Results for the Consumption Function (equation (23))

	ARDL		**CS-ARDL**	
Lags	(1,1)	(2,2)	(1,1)	(2,2)
Regressions with Key variables				
$\widehat{\partial'^{*}}_{Y}$	0.064	0.444***	0.249*	0.341***
	(0.060)	(0.062)	(0.269)	(0.059)
$\widehat{\partial'^{*}}_{RW}$	-0.283	0.224	1.509	0.197
	(0.224)	(0.222)	(1.492)	(0.319)

$\widehat{\partial'^*}_{WS}$	-0.065 (0.232)	0.101 (0.140)	0.419 (0.343)	0.308* (0.184)
$\widehat{\partial'^*}_{LP}$	0.456 (0.692)	0.028 (0.019)	0.144 (0.228)	0.031 (0.019)
$\widehat{\partial'^*}_{DH}$	0.048 (2.076)		0.029 (0.050)	

Refer to Table 5.5 notes

Under the extended regression (Table 5.8) for the short-run consumption function we find positive and significant effects of Y on C – as in Table 5.7 – using the CS-ARDL method. We also observe a positive and statistically significant relationship between WS and C under the CS-ARDL (2,2) specification. The extended regressions estimates reported for RW are generally not significant and, as such, consistent with the baseline short-run dynamics.

VIII. Investment

In this sub-section we estimate a logarithmic formulation of the investment function, following closely the approach adopted in Stockhammer et al. (2008). In that work, the estimated investment equation is derived as an approximate specification of the general Kaleckian double-sided relation between investment and profits (the inverse of the wage share) around a dynamic equilibrium.

This general investment function is specified as:

$$lnI_{it} = f(\,lnY_{it}, lnRW_{it}, lnWS_{it}, lnINT_{it}, lnDH_{it}, lnDB_{it}) + \varepsilon_t \qquad (1)$$

where gross investment, I_{it} is the sum of residential and private investments, excluding government investment. Tables 5.9 to 5.12 report the long-run and short-run coefficient estimates of the explanatory variables in the above investment function. We estimate a baseline regression without DH and DB and add them later in the expanded regression. Long-run estimates from the investment function are summarized in Tables 5.9 and 5.10.

Estimates of Long-run effects

According to Table 5.9, the adjustment coefficients for all three panels ($for\ \rho = 1,2$) have the correct negative sign (except for the CS-ARDL (2,2) result) and are statistically significant, which implies that cointegration exists between the variables. The estimated coefficients for income, female-to-male wage ratio and the wage share have the expected signs across all the different specifications. The size of the long-run estimates of income (Y) corresponds well with the existing literature and the theoretical underpinnings regarding the relationship between aggregate income and investment.

The estimates available from the different specifications also suggest the presence of a negative long-run distributional effect, as significant estimates of the average degree of responsiveness of the wage share for the entire panel of countries range from -0.19 to -0.21 when $\rho = 1$. Our CS-ARDL (1,1) estimate (-0.19) does, however, seem to be much lower than previous estimates available in Stockhammer et al. (2008) and Onaran and Galanis (2012). It should be noted that the estimated numerical coefficient of WS depends on a number of factors that may not have been accounted for in previous estimations – one of which may be the degree of gender equality.

While the main intuition of our framework suggests that an increase in the degree of gender wage equality may have an ambiguous effect on investment, we notice that all ARDL and DL specifications point to a negative effect of RW on investment in the long run in our baseline estimation.[16] However, the interest rate coefficient under the ARDL and DL specifications does produce ambiguous, though not significant, results in terms of sign; we do not find any significant results for the interest rate variable.

Table 5.4:Baseline Regression results for the Investment function (equation (24))

	ARDL		**CS-ARDL**		**CS-DL**
Lags	(1,1)	(2,2)	(1,1)	(2,2)	ρ=1
	Regressions with Key variables				
$\hat{\boldsymbol{\theta}}_{\boldsymbol{Y}}$	1.293***	2.918***	1.059***	1.318***	2.605***
	(0.152)	(0.828)	(0.018)	(0.027)	(0.350)
$\hat{\boldsymbol{\theta}}_{\boldsymbol{RW}}$	-0.388	-3.190	-0.607***	-0.117	-0.078
	(2.102)	(1.560)	(0.126)	(0.096)	(0.894)
$\hat{\boldsymbol{\theta}}_{\boldsymbol{WS}}$	-0.879	-0.704**	-0.194**	-0.037	-0.214*
	(0.630)	(1.352)	(0.128)	(0.096)	(0.271)

[16] See Section 5.4.2.1

$\hat{\theta}_{INT}$	-0.017	0.581	-0.085	0.0004	-0.054
	(0.052)	(0.397)	(0.039)	(0.013)	(0.055)
$\hat{\lambda}$	-0.591***	-0.747***	-0.079***	0.006	
CD	101.48	92.49***	9.78***	4.80***	-0.23
$\vec{\Delta}$	23.787**				
Refer to Table 5.5 notes					

1.1.1.1.2. Robustness to additional explanatory variables

We extend our baseline investment function by including variables previously highlighted as potential determinants of the level of investment. The estimated long-run coefficient of aggregate income remains positive under the ARDL models and the CS-ARDL (1,1) model. However, parameter estimates for the other explanatory variables show very little sensitivity to the baseline results with respect to their benchmark assumptions. We do however find a consistently negative effect of RW and WS on the level of investment when lag $= 2$.

Although, the estimated adjustment coefficients are qualitatively and quantitatively similar, the results in Table 5.10 do not provide any consistent evidence of the effect of changes in household debt and business debt on the level of residential and private investment. Another important result is that, across all specifications, the CD-test statistics large enough that the null hypothesis of cross-sectional independence is strongly rejected at the 1% level in all cases. CS-DL results are omitted from the extended investment regression due to the increased number of regressors which substantially reduces the degrees of freedom even with the suggested truncated lag length of one.

Table 5.5. Extended regression results for the Investment function (equation 24)

	ARDL		CS-ARDL			
Lags	(1,1)	(2,2)	(1,1)	(1,1)	(2,2)	(2,2)
Regressions with Key variables						
$\hat{\theta}_Y$	1.282***	1.724***	0.709	1.209	0.941***	-0.269
	(0.326)	(0.233)	(0.795)	(2.381)	(0.011)	(0.175)

$\hat{\boldsymbol{\theta}}_{RW}$	-1.586 (3.435)	-0.183 (0.347)	3.532 (3.389)	-0.279 (0.436)	-0.093** (0.047)	-0.287*** (0.092)
$\hat{\boldsymbol{\theta}}_{WS}$	1.926 (1.983)	0.053 (0.018)	1.394 (1.333)	0.039 (0.096)	-0.253*** (0.065)	-0.089** (0.043)
$\hat{\boldsymbol{\theta}}_{INT}$	0.029 (0.058)	0.046 (0.455)	0.005 (0.157)	0.041 (0.042)	-0.046*** (0.013)	2.696*** (0.307)
$\hat{\boldsymbol{\theta}}_{DH}$	-0.248 (0.257)	0.067*** (0.005)	1.404 (1.161)	1.209 (2.381)	0.328*** (0.025)	
$\hat{\boldsymbol{\theta}}_{DB}$	-0.121 (0.209)	-0.009 (0.534)		0.045** (0.023)		0.219*** (0.030)
$\hat{\boldsymbol{\lambda}}$	-0.795***	-0.188**	-0.647***	-0.084***	-0.178**	-0.046***
CD	44.62***	81.73***	78.61***	54.55***	16.86***	80.29***

Refer to Table 5.5 notes

Short-run Dynamics

The estimates of the short-run country-specific error correction models provide evidence that suggest that long-run and short-run income effects on investment are mostly similar in the ARDL and CS-ARDL specifications (for the baseline and extended regressions; Table 5.11 and 5.12).

In Table 5.11, using the ARDL approach, we find that a higher level of gender wage equality has a positive [but not statistically significant] effect on investment in the short-run; the opposite is true of the CS-ARDL specifications where we observe negative estimates as with our long-run estimates in Table 5.10 above. We also find that across the baseline results in Table 5.11 and the results on the extended regression in Table 5.12, the estimated short-run coefficients of the labor share of income are largely inconsistent in terms of the nature and direction of its association with aggregate investment. However, only negative coefficients on

WS appear to be statistically significant in the short-run as observed using the CS-ARDL (2,2) in the baseline case and ARDL (1,1) in the extended regression. These results for WS and RW are consistent with our initial assumption of the potentially ambiguous effect of WS and RW on Investment.

Table 5.6: Regression results for the Investment function (equation (24))

	ARDL		CS-ARDL	
Lags	(1,1)	(2,2)	(1,1)	(2,2)
Regressions with Key variables				
$\widehat{\partial'^{*}}_{Y}$	1.417***	2.303*	0.635	1.266*
	(0.227)	(0.045)	(0.757)	(0.734)
$\widehat{\partial'^{*}}_{RW}$	1.171	0.848	-9.323	-0.742
	(0.909)	(1.354)	(7.428)	(0.467)
$\widehat{\partial'^{*}}_{WS}$	0.429	-0.032	-0.647	-1.574**
	(0.563)	(0.413)	(0.802)	(0.790)
$\widehat{\partial'^{*}}_{INT}$	0.017	0.017	0.045	-0.225
	(0.024)	(0.023)	(0.065)	(0.188)
where $\widehat{\partial'^{*}}$ represents the first-lag short-run coefficients related to the respective regressors				

We also find no response of aggregate investment to short-run changes in interest rates, as in our long-run case. This result for the interest rate does seem to be in line with Kalecki's considerations that the interest rate is a less significant factor than the effect of aggregate profitability on the level of investment.

Table 5.7: Regression results for the Investment function (equation (24))

	ARDL		CS-ARDL	
Lags	(1,1)	(2,2)	(1,1)	(1,1)
Regressions with additional variables				
$\widehat{\partial'^{*}}_{Y}$	2.108***	2.184***	1.009***	2.073***

	(0.269)	(0.268)	(0.354)	(0.231)
$\widehat{\partial'^*}_{RW}$	0.407*	0.358	-0.307	0.387
	(0.243)	(0.222)	(1.889)	(0.737)
$\widehat{\partial'^*}_{WS}$	-0.057*	-0.034	2.149	0.193
	(0.137)	(0.142)	(2.235)	(0.374)
$\widehat{\partial'^*}_{INT}$	0.003	0.005	-0.045	0.022
	(0.014)	(0.018)	(0.053)	(0.013)
$\widehat{\partial'^*}_{DH}$	0.144**	0.128	0.117	
	(0.072)	(0.078)	(0.149)	
$\widehat{\partial'^*}_{DB}$	-0.089	-0.128**		0.047
	(0.064)	(0.062)		(0.047)
Refer to Table 5.7 notes				

The picture is somewhat clearer in the case of short-run debt effects. For household debt-to-GDP, we find all coefficients under the different specifications to be positive and very close in magnitude (ranging from 0.117 to 0.144) indicating a potential positive relationship between household debt and [residential] investment. However, as in the long-run case, the level of business debt has an ambiguous effect on investment in the short-run as suggested by the signs and statistical significance of the estimates under the ARDL and CS-ARDL models.

IX. Foreign Sector: Exports

To model the foreign sector, we estimate separate import and export equations as in Stockhammer and Wildauer (2015). In so doing, we are able to model imports as a function of domestic income (Y) and exports (exports are taken to be a function of foreign income (FY)). For our estimations, we employ the following export (X) and import (M) functions:

$$lnX_{it} = f(\, lnRW_{it}, lnWS_{it}, lnEX_{it}, lnFY_{it}) + \varepsilon_t \qquad (2)$$

$$lnM_{it} = f\left(lnY_{it,} lnRW_{it}, lnWS_{it}, lnEX_{it}, lnX_{it}\right) + \varepsilon_t \qquad (3)$$

Estimates of Long-run effects

Table 5.8: Regression results for the Export function (equation (44))

	ARDL		CS-ARDL		CS-DL
Lags	(1,1)	(2,2)	(1,1)	(2,2)	ρ =1
Regressions with Key variables					
$\hat{\boldsymbol{\theta}}_{RW}$	1.603***	1.486***	1.278***	0.689**	1.519***
	(0.444)	(0.433)	(0.290)	(0.331)	(0.006)
$\hat{\boldsymbol{\theta}}_{WS}$	-0.660***	-0.553***	-1.072***	-1.648***	-1.465**
	(0.171)	(0.159)	(0.238)	(0.290)	(0.407)
$\hat{\boldsymbol{\theta}}_{EX}$	-0.217***	-0.193***	-0.507***	-0.131*	-0.538**
	(0.068)	(0.063)	(0.076)	(0.071)	(0.228)
$\hat{\boldsymbol{\lambda}}$	-0.084***	-0.046***	-0.235***	-0.315***	
CD	84.00***	92.49***	12.44***	8.67**	2.02**
$\vec{\Delta}$	9.008*				
Refer to Table 5.5 notes					

The results of the export function regression in Tables 5.13 and 5.14 indicate that the error-correction coefficients fall within the dynamically stable range (being statistically significant and negative), and therefore the null hypothesis of no long-run relation is rejected. This finding indicates that there is compelling evidence for conditional convergence to country-specific dynamic equilibriums in our sample of 45 countries.

Furthermore, results from Table 5.13 suggest that in the long run, the wage share (WS) and exchange rates (EX) are negatively associated with the level of exports, as expected. However, we observe significant positive effects of the female-to-male wage ratio on total exports for the entire panel with coefficients ranging from 0.6 to 1.6, depending on the methodology employed. Also, we find evidence of cross-sectional dependence under all specifications – although the CD statistics appear more marginal for the CS estimates. It is also worth noting that the reported results in Table 5.13 are consistent across the different model specifications. More importantly, these findings for wage share and exchange rates follow our *a priori* expectations and are also in line with the results of Stockhammer and Wildauer (2015) and Onaran and Obst (2016). The results for RW suggest that increasing women's wages may positively affect the level of exports. We do however expect that this result may be different in the short-run.

1.1.1.1.3. Robustness to additional explanatory variables

The results of the robustness check confirm the importance of the degree of gender wage equality and exchange rates as drivers of long-run changes in exports in our panel; the long-run gender effect in the baseline and the expanded regression specifications all prove significant under our preferred CS-ARDL model, corroborating the results of Seguino (1997). On the other hand, the results related to the wage share are much more ambiguous. The negative long-run wage share effect suggested by the results from the baseline regression are only consistent with results in the extended model when one lag is imposed. We find positive but insignificant coefficients in the various specifications of the extended regression when $\rho = 2$. With respect to the long-run effect of foreign income (FY) on exports, we find a mostly positive and statistically significant relationship between FY and exports following our *a priori* expectations.

Table 5.9: Regression results for the Export function (equation (44))

	ARDL		**CS-ARDL**		**CS-DL**
Lags	(1,1)	(2,2)	(1,1)	(2,2)	ρ=1
Regressions with additional variables					
$\hat{\boldsymbol{\theta}}_{RW}$	0.357	0.363	1.136***	0.503**	1.942
	(2.755)	(2.125)	(0.297)	(0.172)	(0.997)
$\hat{\boldsymbol{\theta}}_{WS}$	-1.509	0.688	-0.933*	0.543	0.455
	(2.306)	(0.889)	(0.478)	(0.290)	(0.892)
$\hat{\boldsymbol{\theta}}_{EX}$	-0.048	-0.044	-2.158***	-0.961***	-0.653*
	(0.191)	(0.105)	(0.211)	(0.119)	(0.336)
$\hat{\boldsymbol{\theta}}_{FY}$	2.256***	-0.037	0.351**	1.083***	2.091***
	(0.336)	(0.346)	(0.137)	(0.071)	(0.512)

$\hat{\lambda}$	-0.549***	-0.043**	-0.037***	-0.091***	
CD	93.13***	24.30**	50.9***	68.03***	9.75***
Refer to Table 5.5 notes					

Short-run Dynamics

The short-run dynamics for the export function are reported in Tables 5.15 and 5.16. In both reports, we find no significant short-run association between the female-wage ratio and the level of exports. We ascribe the lack of a short-run reaction of exports to changes in RW primarily to the dominance of developed countries in our panel sample, as we expect that the level of exports in developing countries responds more swiftly to changes in the level of gender wage equality in comparison to more developed economies.

Table 5.10: Regression results for the Export function (equation (44))

	ARDL		**CS-ARDL**	
Lags	(1,1)	(2,2)	(1,1)	(2,2)
Regressions with Key variables				
$\widehat{\partial'^{*}}_{RW}$	0.036	0.050	-0.469	-0.462
	(0.095)	(0.108)	(0.808)	(0.775)
$\widehat{\partial'^{*}}_{WS}$	-0.239***	-0.226**	-0.449*	-0.075
	(0.074)	(0.075)	(0.5110	(0.469)
∂'^{*}_{EX}	-0.082**	-0.083**	-0.054	-0.250**
	(0.037)	(0.037)	(0.065)	(0.093)
Refer to Table 5.7 notes				

Table 5.11: Regression results for the Export function (equation (44))

	ARDL		**CS-ARDL**	
Lags	(1,1)	(2,2)	(1,1)	(2,2)
Regressions with additional variables				
$\widehat{\partial'^{*}}_{RW}$	0.452	-0.404	0.013	-1.924
	(1.550)	(1.119)	(0.681)	(1.685)

$\widehat{\partial'^*}_{WS}$	0.243	0.008	1.699	0.603
	(1.149)	(0.421)	(1.292)	(0.585)
$\widehat{\partial'^*}_{EX}$	-0.049	-0.069	0.103	-0.009
	(0.064)	(0.073)	(0.249)	(0.072)
∂'^*_{FY}	1.192***	2.199***	0.411	-0.073
	(0.197)	(0.399)	(0.855)	(0.224)
Refer to Table 5.7 notes				

X. Foreign Sector: Imports

Estimates of Long-run effects

Tables 5.17 and 5.18 presents the estimated results of the import equation for the entire panel. In summary, the impact of domestic income on the level of import is positive and statistically significant in most cases in the baseline and extended regressions. Likewise, the export coefficient in Table 5.18 suggests a positive and [mostly] significant impact on the level of imports for our panel.

Table 5.12: Regression results for the Imports function (equation (45))

	ARDL		CS-ARDL		CS-DL
Lags	(1,1)	(2,2)	(1,1)	(2,2)	ρ =1
Regressions with Key variables					
$\hat{\theta}_{Y}$	2.094***	1.543***	1.638	1.493***	1.608***
	(0.277)	(0.035)	(1.450)	(0.132)	(0.296)
$\hat{\theta}_{RW}$	-2.252	1.201***	0.461***	3.722	-1.165
	(1.444)	(0.105)	(0.144)	(1.045)	(1.835)
$\hat{\theta}_{WS}$	1.109**	0.628***	-1.104	-0.049	-1.214
	(0.564)	(0.067)	(1.864)	(0.121)	(2.221)
$\hat{\theta}_{EX}$	0.167	0.107**	-1.186	-0.040	-0.062**
	(0.208)	(0.048)	(1.174)	(0.049)	(0.138)
$\hat{\lambda}$	-0.595***	-0.216***	-0.709***	-0.031**	
CD	67.22	84.33***	12.45***	19.12***	1.35**

$\vec{\Delta}$	98.766***

Refer to Table 5.5 notes

Table 5.13:Extended regression results for the Imports function (equation (45))

	ARDL		CS-ARDL		CS-DL
Lags	(1,1)	(2,2)	(1,1)	(2,2)	ρ=1
Regressions with additional variables					
$\hat{\theta}_Y$	0.365	0.576***	1.607***	1.137***	0.931
	(0.451)	(0.053)	(0.477)	(0.147)	(0.751)
$\hat{\theta}_{RW}$	-1.751	0.359***	9.056	0.349**	1.387
	(1.304)	(0.089)	(8.805)	(0.167)	(1.790)
$\hat{\theta}_{WS}$	0.036	-0.432***	3.099***	0.224**	2.207
	(0.522)	(0.082)	(0.954)	(0.087)	(1.597)
$\hat{\theta}_{EX}$	0.795	0.291***	0.481*	0.024	-0.566
	(0.615)	(0.042)	(0.253)	(0.038)	(0.366)
$\hat{\theta}_X$	0.730***	0.581***	0.294	0.448***	0.979*
	(0.151)	(0.032)	(0.278)	(0.069)	(0.510)
$\hat{\lambda}$	-0.739***	-0.190***	-0.877***	-0.141**	
CD	61.93	34.82***	87.84	18.46***	1.77*

Refer to Table 5.5 notes

However, the results for the effect of the wage share and exchange rates are ambiguous as we observe statistically significant effects in opposite directions in both tables. For the female-to-male wage ratio, we find our estimated coefficients to be significant only when pointing to a positive relationship between RW and imports. The theoretical literature on this issue is also, by and large, inconclusive but we do expect clearer inference from the results for our different country groups.

Short-run Dynamics

Our short-run findings for our import function are closely similar in magnitude and direction to our long-run results. For the sake of brevity, we find a positive effect of domestic income and exchange rates on imports, following our benchmark assumptions. We also find, as in our long-run case, that the wage share and ratio of female-male wages has no clearly identified direction of association with the level of imports for all 45 countries in our panel.

Table 5.14: Regression results for the Imports function (equation (45))

	ARDL		CS-ARDL	
Lags	(1,1)	(2,2)	(1,1)	(2,2)
Regressions with Key variables				
$\widehat{\partial'^*}_Y$	0.803***	1.506***	0.814*	1.405***
	(0.170)	(0.165)	(0.439)	(0.152)
$\widehat{\partial'^*}_{RW}$	1.614	-0.882	0.330**	-1.987
	(1.213)	(1.228)	(0.129)	(2.658)
$\widehat{\partial'^*}_{WS}$	0.022	0.953	-0.866	-0.015
	(0.178)	(0.801)	(0.999)	(0.104)
$\widehat{\partial'^*}_{EX}$	0.106	0.109	-0.561	-0.026
	(0.099)	(0.100)	(0.656)	(0.039)
Refer to Table 5.5 notes				

Table 5.15: Extended regression results for the Imports function (equation (25))

	ARDL		CS-ARDL	
Lags	(1,1)	(2,2)	(1,1)	(2,2)
Regressions with additional variables				
$\widehat{\partial'^*}_Y$	0.626***	0.998***	-0.523	1.211***
	(0.176)	(0.187)	(0.811)	(0.186)

$\widehat{\partial'^*}_{RW}$	1.859	1.240	-1.993	0.342**
	(1.199)	(1.259)	(1.387)	(0.144)
$\widehat{\partial'^*}_{WS}$	-0.192	0.525*	1.035	-0.036
	(0.341)	(0.291)	(1.516)	(0.079)
$\widehat{\partial'^*}_{EX}$	0.118	0.169	-0.085	0.025
	(0.252)	(0.113)	(0.267)	(0.042)
$\widehat{\partial'^*}_{X}$	0.108	0.306***	0.069	0.376***
	(0.084)	(0.072)	(0.283)	(0.07

Refer to Table 5.5 notes

XI. Total Effects: Demand-Led Regimes

Using our annual panel dataset, we employed our preferred CS-ARDL approach to account for endogeneity, cross-country heterogeneity, and cross-sectional dependence which arise from unobserved common factors. Our main findings suggest a general long-run cointegrating relationship between the relevant explanatory and dependent variables of the various equations of the aggregate demand components. Furthermore, these results from our formal statistical analyses enable us to infer the long-run and short-run elasticities of the regressors in each of the equations for the aggregate expenditure components.

Our previous estimates concentrate on the impact of income, gender wage inequality and the labor share of income on consumption, investment and net exports. However, to examine the distributional effects on aggregate demand, we need to determine, from our estimated elasticities, the partial effects of the wage share on the various components of aggregate demand. These converted marginal effects can then be used to account for the cumulative effects of the wage share on aggregate demand.[17] Note that, the entire panel, the overall estimates derived from the econometric application is employed in the calculated of

[17] Elasticities are converted into marginal effects using:

$$\frac{\partial Y}{\partial WS} = \hat{\beta}_{C,WS}\left(\frac{C}{WS}\right) + \hat{\beta}_{I,WS}\left(\frac{I}{WS}\right) + \hat{\beta}_{X,WS}\left(\frac{X}{WS}\right) - \hat{\beta}_{M,WS}\left(\frac{M}{WS}\right)$$

the marginal effects. On the hand, for the different country groups, the marginal effects are calculated using the GDP-weighted averages of the estimates for each of the individual countries in that specific sample group. Table 5.21 reports the long-run and short-run marginal effects for our entire panel. Marginal effects and cumulative effects are also reported for our different country groups using the conversions from our CS-ARDL (1,1) model.

Table 5.16: Marginal Effects of 1% change in WS on private excess demand for the 45-country panel.

X	1 lag; 3 CS lags		2 lags; 2 CS lags	
	LR	SR	LR	SR
Consumption	0.671	0.133	1.129	-0.199
Investment	-0.049	-0.163	-0.009	-0.397
Net Exports	-0.11	-0.016	-0.288	-0.014
$\mathbf{Y^{PED}}$	0.512	-0.046	0.832	-0.610
Multiplier	4.484	1.156	4.149	1.603
Total Effect	2.296	-0.053	3.452	-0.978
Long-run: Wage-led growth || Short-run: Profit-led growth				

Table 5.22 reports the marginal effects of the wage share on the components of aggregate demand and its cumulative effect on aggregate demand based on average parameters. Our results indicate overall long-run wage-led growth and short-run profit-led growth following the respective signs of private excess demand (Y^{PED}).[18] This finding is in line with the results of Sánchez and Luna (2014) for Mexico and is consistent with Blecker's (2016) argument that the magnitude of the impact of the wage and profit share on the components of aggregate demand may depend substantially on the time period examined, which in turn may explain

[18] Private excess demand as defined by Onaran and Galanis (2012) is the sum of the partial effects of the components of demand prior to the multiplier effects.

the conflicting results present in the literature on the impact of the distribution on aggregate demand:

"Some distributional effects may be more important in the short run (over a few quarters or years, or the length of an ordinary business cycle), while others are likely to be more important in the long run (across, say, one or more decades). In particular, … the positive effects of higher profit shares (or lower labor costs) on investment and net exports are mainly short-run phenomena, while the sensitivity of workers' consumption to their wage income is, if anything, likely to be stronger in the long run." (Blecker, 2016 p. 3)

Furthermore, we calculate our long-run and short-run multipliers using equation (27) above. Our results suggest that the long-run multipliers (accumulated effects) are substantially larger than our reported interim (short-run) multipliers. We note that our short-run multipliers are consistent with what is usually observed in the literature (Onaran and Galanis, 2012; Keifer and Rada, 2014). With regards to the long-run multipliers, we have no research to compare our findings to, but it is however important to note that our relevant long-run coefficients (elasticities) appear statistically significant more frequently than do our short-run results.[19] Finally, the last row of Table 5.21 shows the total effect on aggregate demand from a change in income distribution when the multiplier mechanism is accounted for (see equation (27) i.e. g_2). We find, as expected, a larger impact on equilibrium income from changes in the wage and profit share when we introduce the multiplier.

Subsample Cumulative Effects

Evidence for our panel suggests a long-run and short-run relationship between functional income distribution and aggregate demand with demand-led growth observed to be profit-led in the short-run and wage-led in the long-run. However, this result for demand-led growth may vary for certain country groups. In what follows, we report the marginal and cumulative effects for our different country groups. As stated in earlier sections, the 46 countries in the panel are grouped into high-income, middle-income/SICs and low-income/LIAEs countries. Marginal effects are calculated using the same formula as in our panel case; however, the elasticities employed in the calculation are derived from taking the average of the country-

[19] We employ both statistically significant and non-significant point estimates in the calculation of our partial and cumulative effects due to the scant evidence of statistical significance of our short-run estimates. However, this lack of statistical significance of some of these estimates implies a lack of precision in the estimation of the partial effects.
Furthermore, the wider confidence intervals of our parameters – in relation to the estimates – may indicate instability; thus, we are uncertain that the value of the corresponding parameter in the underlying regression model is zero.

specific mean group estimates for the countries in each income group. Similarly, in this case, the consumption share is also derived from the sub-sample weighted sum. These sub-sample averages are calculated using the GDP-weighted averages of the various aggregate demand components (C, I, X, M) normalised by income and the GDP-weighted average of the wage share.[20]

Table 5.17: Long-run WS effects on Aggregate demand for different Income groups

	PANEL	DEVELOPED COUNTRIES	SICs	LIAEs
Consumption	0.671	-0.661	2.879	4.307
Investment	-0.049	3.701	-0.274	0.169
Net Export	-0.11	-0.493	0.498	-0.154
$\mathbf{Y}^{PED}$	0.512	2.547	3.103	4.322
Multiplier	4.484	0.561	1.769	0.563
Total Effect	2.296	1.429	1.489	2.433
Openness	7%	21%	36%	61%

Tables 5.22 and 5.23 report the results where we present only the partial, cumulative and multiplier effects for the long and short-run cases using the estimates from our preferred CS-ARDL (1,1) model. Notably, the error correction coefficient is negative and significant across all regressions, indicating cointegration between our regressors and their respective AD components.

Several studies point to the importance of economic structure and openness to trade in determining the nexus between income distribution and aggregate income (see, for example, Blecker (2016) and Stockhammer et al. (2011)). In the same vein, Onaran and Galanis (2012) show that economies with a higher degree of openness are more likely to be profit-led than wage-led – at least in the short-run. This is perhaps because net export effects can be directly

[20] The cumulative effect for each country group is derived thus:

$$\frac{\partial Y}{\partial WS} = \hat{\beta}_{C,WS}\left(\emptyset\frac{C}{Y}\right)\frac{1}{\emptyset WS} + \hat{\beta}_{I,WS}\left(\emptyset\frac{I}{Y}\right)\frac{1}{\emptyset WS} + \hat{\beta}_{X,WS}\left(\emptyset\frac{X}{Y}\right)\frac{1}{\emptyset WS} + \hat{\beta}_{M,WS}\left(\emptyset\frac{M}{Y}\right)\frac{1}{\emptyset WS}$$

where $\emptyset$ represent the income weight for each country within a respective income group. As in the previous marginal effects conversion, the elasticities are converted to marginal effects and then normalised by income as in Stockhammer and Wildauer (2015). However, for the various country groups, these elasticities are calculated as GDP-weighted averages for the countries in each sub-panel.

associated with the openness index (i.e., the share of exports and imports in GDP), such that small open economies, for example, may be more sensitive to net export volatility thus potentially leading to circumstances where negative effects of a rise in the wage share are enough to overshadow any positive wage share effects on investment and consumption. This, we find to be the case in the short-run for our low-income countries (see Table 5.23).

The results in Table 5.22 show the long-run effects for our different country groups. Our results suggest that growth is wage-led for all country groups and the entire panel. However, there are several interesting patterns. First, the evidence seems to suggest a negative wage share effect on net exports for the high and low-income countries. While the results for the high-income countries follow the postulation that large and relatively closed economies (which are more likely to be high-income countries) have a small net export effect compared to their consumption effect and, as such, are strongly wage-led. We find that the net export effect in the LIAEs is relatively smaller compared to the other country groups and more importantly, smaller relative to the consumption effect – hence confirming the observed long-run wage-led growth regime. Secondly, following from our first observation, the degree of openness may not always be directly linked with the demand regime of an economy when the level of development and long-run time dimension comes into play. Third, the wage share effect on consumption for high-income countries is surprisingly negative overall and becomes larger and positive with lower levels of aggregate income (i.e., SICs and LIAEs).

We also find a considerably large positive effect on investment from changes in the wage share for the high-income countries in our panel. This large positive effect is what determines the wage-led regime for this country group and not the wage effect on consumption as would be expected. In fact, we find a negative consumption effect for this group of countries. This relationship may be explained by the argument that potentially rising wages can serve as a boost to the long-term profitably of firms by stimulating investment (Blecker, 2016). However, there is no comprehensive answer to the question of why corporate investment and profitability will be positively affected by a higher wage share, but the positive impact of increasing wages on residential investment is a possibility.

Looked at from this angle, one could argue that growth in the wage share will be necessary to encourage household residential investment expenditure and, in a similar vein, also encourage producers to undertake investment. This sustained investment should in turn stimulate positive changes in production that underpin productivity growth. We find a similar positive, albeit smaller, effect of the wage share on investment for our low-income countries.

On the other hand, we find a negative wage share effect on investment for the middle-income countries (SICs) in our panel. This may suggest that an increase in the wage share may indicate a loss of profit to capitalists or businesses – which may in turn discourage future investment.

Table 5.18: Short-run effects on aggregate demand for different income groups

	PANEL	DEVELOPED COUNTRIES	SICs	LIAEs
Consumption	0.133	0.076	0.131	1.333
Investment	-0.163	0.453	-0.548	2.123
Net Export	-0.016	-0.119	-0.348	-4.004
Y^{PED}	-0.046	0.410	-0.765	-0.548
Multiplier	1.156	1.675	0.718	1.783
Total Effect	-0.053	0.687	-0.549	-0.977

Our short-run results – reported in Table 5.23 – appear notably different from their long-run counterparts. Here, we find evidence to support a profit-led demand regime for the entire panel. This profit-led growth hypothesis is also confirmed for our middle-income (SICs) and low-income (LIAEs) countries. For the SICs, we find negative wage share effects on investment and net exports which overshadow the positive WS effect on consumption – hence a profit-led demand growth regime. We find the co-existence of a long and short-run wage-led growth regime only for our high-income countries' group.

Table 5.19: Growth Regimes Summarised

	PANEL	DC	SICS	LIAES
Long-run	Wage-led	Wage-led	Wage-led	Wage-led
Short-run	Profit-led	Wage-led	Profit-led	Profit-led

Moreover, these findings follow the hypothesis that the degree of openness of an economy directly corresponds to the net export effects. Following the benchmark assumptions, we find that relatively closed economies (these are often larger developed countries with the larger share of global production) are more strongly wage-led, while medium-sized open economies (such as small SICs) tend to have relatively smaller net export effects leading them to often follow a profit-led growth regime. On the other hand, low-income economies, which are often small open economies, as in our panel, tend to have large negative net export effects (-4.004) which may result in the total demand regime becoming profit-led. These *a priori* expectations of a corresponding relationship between openness and the external sector seem to hold only for our short-run demand-led growth determination. Furthermore, the short-run multipliers, as earlier mentioned, are more consistent with the existing literature – these are especially similar to the multipliers derived in Onaran and Galanis (2012).

XII. Gender Effects on The Macroeconomy

Having identified the growth regimes of our different country groups above, we now estimate the relative contributions of gender inequality on each of our macroeconomic aggregates. In doing so, we also account for the role of economic structure in affecting the relationship between gender equality and macroeconomic outcomes. An obvious way to employ gender as an explanatory macroeconomic tool is to disaggregate at least one of the components of aggregate demand by gender. As previously described, such a disaggregation fits well with the Bhaduri and Marglin (1990) demand-led model, which emphasizes the distributional differentiation of economic agents.

To determine the macroeconomic effects of gender inequality, we investigate under what conditions the different country groups – and the entire panel – display a gender equality-led growth regime in the long or short run. As earlier explained, a gender equality-led growth regime is one that is 'gender cooperative' in which an increase in gender equality results in injections exceeding leakages ($S + M < I + X$) i.e., a redistribution stimulates aggregate demand, leading to an increase in output (as in Seguino (2012)). Tables 5.25 and 5.26 report the long and short-run RW elasticities for each of the individual AD components using our preferred CS-ARDL (1,1) model.[21] At the panel level, the gender effects are quite modest and expansionary in the long-run but larger and contractionary in the short-run. These magnitudes of effects are as expected as gender effects are more likely to persist strongly in the short-run than in the long-run.

Table 5.20: Summary of Long-run Gender Effects

[21] The calculated elasticities are employed whether or not they appear statistically significant, in this part of the analysis, as in the demand-led regime determination above.

L.R	PANEL	DC	SICS	LIAES
S	0.008	10.23	-0.377	6.743
M	0.461	-1.875	0.632	0.362
SUM	0.469	8.355	0.255	7.105
I	-0.607	1.826	-0.388	1.723
X	1.278	1.884	4.612	2.387
SUM	0.671	3.71	4.224	4.11
EFFECT	E	C	E	C

Note: where E refers to an Expansionary effect and C refers to a Contractionary effect

Table 5.25 presents the long-run directional effects of greater gender equality on the various components of aggregate demand. While our results for the entire panel suggest that greater gender equality exerts expansionary pressures on economic growth, this inference does not hold firmly when viewed at the different stages of economic development. Particularly, we find that greater gender equality imposes contractionary pressures on growth for developed and low-income countries, while it produces expansionary pressure on SICs/middle-income countries.

Furthermore, Table 5.26 provides a summary of how lower gender wage gaps affect various macroeconomic outcomes in the short-run, and the implications for growth. As might be expected, in the short-run, greater gender inequality is contractionary for the entire panel due to the potential negative effects of higher female wages on competitiveness. As in our long-run case, we observe varying results for the different country groups; particularly, the short-run impact of greater gender equality on the process of economic growth is contractionary for the developed and semi-industrialized countries but expansionary for the low-income countries.

A few things stand out from our results. First, we observe that the magnitude of the gender effect on savings and imports are much larger for LIAEs and developed economies both in the long and short-run. However, while these gender effects are large enough to compel a contractionary squeeze on growth in developed economies in the long and short run, in LIAEs they are large enough to drive contractionary pressures on growth only in the long run.

Table 5.21: Summary of Short-run Gender Effects

S.R	PANEL	DC	SICS	LIAES
S	0.652	-1.446	-0.076	0.664
M	0.330	-0.317	-0.076	0.435
SUM	0.982	-1.129	-0.152	1.099
I	-9.323	0.534	-0.236	-0.636
X	-0.469	-0.664	-0.580	3.576
SUM	-9.792	-0.13	-0.816	2.94
EFFECT	C	E	C	E

Secondly, our results further underscore the crucial role played by gender equality in the relationship between equality and macroeconomic outcomes. As observed, gender wage equality appears to have differential effects in the long and short run depending on the nature of economic development and economic structure. For example, we find that in the short run, gender-based wage differentials are only linked to export growth in high-income and low-income countries; however, the processes by which this occurs in both country groups are entirely different. In the high-income group (DC), we observe that this expansionary stimulus to short-run growth is mostly a result of the positive impact of greater gender wage equality on investment. However, the exact opposite is true for low-income countries in the short-run as the potential positive push on growth is due to the positive impact of greater gender equality on consumption, imports and exports where, in this case, higher female wages negatively impact investment. Furthermore, in the long-run our results suggest that greater gender wage equality produces a contractionary stimulus on growth for both high-income and low-income countries, implying directly opposite aggregate demand effects in the long and short-run for low and high-income countries. While these long-run effects are not unexpected for the low-income countries as they advance towards higher development, we are unsure as to the reasoning behind the potential contractionary pressures on growth due to sustained increases in female wages in high-income countries.

Table 5.22: Summary of Gender Effects

	PANEL	DC	SICS	LIAES
Long-run Effect	Gender Cooperative	Gender Conflictive	Gender Cooperative	Gender Conflictive
Short-run Effect	Gender Conflictive	Gender Conflictive	Gender Conflictive	Gender Cooperative

In summary, our short-run results follow our *a priori* expectations. As stated in earlier sections, we expect that, due to the adoption of export orientation of many middle-income or semi-industrialised economies, low female wages may indeed act as a stimulus to growth. A large body of literature has linked gender employment and wage discrimination, embedded in social and labour market practices to increased export competitiveness which are largely dependent on low unit labour costs, which women can readily provide under these conditions (Nash and
Fernández-Kelly, 1983; Deyo, 1989; Hsiung, 1996; Seguino, 1997; Blecker and Seguino, 2002). In addition, we also expect greater gender wage equality to improve demand-led growth conditions in LIAEs and DCs. In LIAEs, it is assumed that women are more concentrated in subsistence farming and domestic unpaid care work. As such, any improvement in female wages will lead to increased productivity in the agricultural and care sector as well as in female labour force participation, all of which work through increasing productivity and the human capital endowments of women and children (Darity, 1995; Seguino, 2012).

In contrast, our long-run results do not necessarily conform to any previous expectations. The most puzzling aspect of our long-run results is that they are exactly opposite to the observed short-run results for all the country groups and for the entire panel. While growth can be assumed to respond to expansionary pressures in the long-run for the entire panel, only SICs can be assumed to follow a similar pattern. However, contrary to our expectations, our results suggest that this expansionary pressure on growth is mostly mobilised by the positive effect of a sustained increase in female wages on exports in the long-run. This striking result may imply that while gender inequality is linked to export-led growth at a country's adoption stage of export orientation, a continual closing of the gender gap may in the long-term result in increased export potentials due to the implications on labor productivity transmitted through rising female educational attainment, the diversification of the female force and increased female bargaining power due to increases in their discretionary income (see Seguino, 2013). This result may buttress the neoclassical theory which posits that at long-run full employment, women's wages are expected to rise to indicate their increased productivity.[22]

XIII. Growth Contributions

To investigate the potential of gender equality as a driver of growth in the period examined (1985 – 2015), we analyze the extent to which gender equality is able to explain changes in the growth of aggregate income by summarising the magnitude of the responses of the aggregate demand components to changes in the female-male wage ratio. To do this, we convert the long-run and short-run output elasticities of the macro aggregates with respect to RW; using these RW elasticities, we then decompose the growth rate of consumption, investment, imports and exports due to gender wage equality over time.

These growth contribution estimates are conducted for all three country groups, applying the formula: $\hat{\beta}_{RW} \Delta RW$. where $\hat{\beta}_{RW}$ is the estimated elasticity using our CS-ARDL (1,1) models for each of the component estimations (C, I, X and M). ΔRW represents the [GDP-weighted] year-on-year change in the female-male wage ratio. The long and short-run results are presented in Tables 5.28 and 5.29 respectively.

The relative long-run growth contributions of gender wage equality to the aggregate demand components above is negative for the entire panel, except for investment. However, when we consider these contributions in terms of economic structure, we observe, in some cases, dissimilar effects. For example, for developed countries we find a negative growth effect of gender wage equality on only aggregate household consumption and imports; the implication is immediately clear on the import side as we can posit that as women's wages in developed economies increase, there is a greater incentive to consume more locally made products as income rises.

Table 5.23: Long-run Growth Contributions of Gender Equality

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[22] Recall that in our typical usage of the "long-run" in this study, we refer to the quasi-equilibrium which implies a succession of many short-run periods. Also, in our reference to growth effects, we are concerned with the potential pressures on – or against – economic growth as is implied in the Kaleckian demand-led growth framework.

	PANEL	DEVELOPED COUNTRIES	SICs	LIAEs
$\hat{\beta}_{C,RW}\Delta RW$	-0.382	-2.320	-0.529	6.602
$\hat{\beta}_{I,RW}\Delta RW$	0.234	0.459	0.149	-1.981
$\hat{\beta}_{X,RW}\Delta RW$	-0.492	0.475	-1.775	-2.744
$\hat{\beta}_{M,RW}\Delta RW$	-0.177	-0.472	-0.243	-0.416**
$\mathbf{Y}^{PED}$	0.512	2.547	3.103	4.322

Contrary to our *a priori* expectations, our results suggest that an upward convergence of female to male wages may result in the overall growth of the savings rate in high-income and middle-income countries. More puzzling is the considerably large and positive growth effect of greater gender wage equality on aggregate household consumption in LIAEs. While the gender effect on the savings rate for the SICs follows the results of Seguino and Floro (2002), our findings for the LIAEs are somewhat contrary to the literature which surmises that higher unpaid work burden of women in the household raises the savings rate (Ertürk and Çağatay, 1995; van Stanvaren, 2002). Therefore, a rise in female wage income in LIAEs should raise the savings rate given that women in these countries take up a disproportionately large share of unpaid care labour, relative to men in SICs and developed economies.

Table 5.24: Short-run Growth Contributions of Gender Equality

	PANEL	DEVELOPED COUNTRIES	SICs	LIAEs
$\hat{\beta}_{C,RW}\Delta RW$	-0.134	-0.056	-0.414	-0.386
$\hat{\beta}_{I,RW}\Delta RW$	3.588	0.134	0.908	0.731
$\hat{\beta}_{X,RW}\Delta RW$	0.181	-0.167	0.223	-4.111
$\hat{\beta}_{M,RW}\Delta RW$	-0.127	-0.079	0.292	-0.500
$\mathbf{Y}^{PED}$	0.512	2.547	3.103	4.322

1.7. CONCLUSION

The discourse on gender equality is one that has become increasingly popular in political and economic spheres – and in many other academic disciplines. A large portion of the debate surrounding gender equality appeals to the ideals of human dignity, fairness, justice, maximum societal well-being and economic growth. A second strand focuses on a discussion of the evidence-based instrumental rationale to inspect the impact of gender [in]equality on economic performance and development. This study focuses on the second set of interests.

Our findings suggest that growth is profit-led in the short-run and wage-led in the long-run. However, we find wage-led growth in the long and short run for high-income countries. In addition, following the estimations of gender effects, we are able to posit that global economic growth is gender equality-led and wage-led in the long-run. This is also the case for middle-income countries. Furthermore, short-run improvements in gender wage equality are consistent with profit-led growth in low-income countries, this result suggests that higher gender equality makes the growth regime for LIAEs more wage-led or less profit-led in the long run.

Furthermore, our results are able to highlight the role of distribution in determining (private domestic) demand and the effect this may have on employment and growth. For example, given that global growth is wage-led, even for countries at different stages of economic development, we can assume that wage moderation is unlikely to stimulate employment in the long-run.

Our findings also suggest that a redistribution in favour of wages may create a path to equality-led growth, which is a core component of sustainable growth, due to its increased effect on demand and technological progress. Such a path to sustainable growth will also need a rebalancing in the form of gender equality which – as evident in the literature – promotes productivity and long-run growth and drives a wage-led growth agenda. Such a wage-led growth strategy which aims at an economic regime that highlights increased demand and innovation as a crucial goal will require strong policy coordination. Further research is needed to fully understand why increasing female wages elicit different responses from macroeconomic aggregates in the long and short run.

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Table 1. Results of Cross-Sectional Dependence Test

Variables	***CD_p Statistic***
Y	161.7164***
C	164.5973***
I	130.0207***
X	146.4400***
M	153.5564***
WS	28.31246***
RW	81.78625***
INT	94.31415***
EX	13.13568**
LP	33.65240***
DH	21.5856***
DB	86.9673***
FY	9.3910**

*****, ** and * denote 1%, 5% and 10% significance levels respectively.**

APPENDIX

Table A1. Country Classification

HIGH INCOME COUNTRIES		**MIDDLE-INCOME COUNTRIES**		**LOW-INCOME COUNTRIES**	
Country	Abbr.	Country	Abbr.	Country	Abbr.
Australia	AUS	Argentina	ARG	Belize	BLZ
Austria	AUT	Botswana	BWA	Burkina Faso	BFA
Belgium	BEL	Brazil	BRA	Egypt	EGY
Canada	CAN	Bulgaria	BGR	Gambia	GMB

Chile	CHL	Colombia	COL	Philippines	PHL
Denmark	DNK	India	IND	Senegal	SEN
Finland	FIN	Macedonia	MKD	Sri Lanka	LKA
France	FRA	Mauritius	MAU	Thailand	THA
Germany	DEU	Mexico	MEX	Tunisia	TUN
Greece	GRC	Peru	PER	Uganda	UGA
Ireland	IRL	Poland	POL		
Italy	ITA	Turkey	TUK		
Japan	JPN				
Korea	KOR				
Macao, China	MAC				
Netherlands	NLD				
New Zealand	NZD				
Norway	NOR				
Singapore	SGP				
Spain	ESP				
Sweden	SWE				
Switzerland	CHE				
United Kingdom	GBR				
United States	US				

Table Axxx. Panel Unit Root Test Results

	Deterministic Trend	LLC	IPS	CIPS
Level				
Y	Trend, Intercept	1.855	13.902	-1.449**
C	Trend, Intercept	-1.085	-0.648	-1.129***
I	Intercept	4.798	8.295	3.946
X	Intercept	1.845	13.426	11.370
M	Trend, Intercept	-1.470*	0.711	-2.442**
WS	Trend, Intercept	-0.913***	-0.986	-0.570
RW	Trend, Intercept	-0.997	-2.130**	-3.556***
INT	Intercept	-16.682***	-8.285***	-0.963
EX	Intercept	-23.433***	-23.030***	-11.928***
LP	Trend, Intercept	-3.552***	-7.372***	-8.021**
DH	Intercept	1.738	11.836	1.092
DB	Trend, Intercept	4.679	9.002	-3.453***
FY	Trend, Intercept	-17.081***	-11.627***	-3.442*
First Difference				
ΔY	Trend, Intercept	-16.642***	-18.367***	-1.423***
ΔC	Intercept	-21.591***	-22.289***	-8.751***
ΔI	Intercept	-24.389***	-24.499***	-9.741***
ΔX	Intercept	-21.780***	-22.663***	-6.269***
ΔM	Intercept	-22.511***	-22.532***	-4.284***
ΔWS	Intercept	-14.979***	-18.650***	-1.275***
ΔRW	Intercept	-17.087***	-20.596***	-3.051**
ΔDH	Intercept	-22.098***	-11.534***	-2.109***

ΔDB	Intercept	-31.342***	-33.453***	-12.098***
Notes:	We determine the optimal lag length using the Schwarz Information Criterion (SIC). Δ is the first difference operator. ***, ** and * denote rejection of H_0 at 1%, 5% and 10% significance levels respectively. The inclusion of a trend term is dependent on the observable characteristics of the series.			