SMU ECONOMICS & STATISTICS
WORKING PAPER SERIES

Skilled & Unskilled Wages in Globalizing World, 1968 - 1998

Davin Chor

May 2001

Paper No. 06-2008

Skilled and Unskilled Wages in a Globalizing World, 1968-1998

Han Ping Davin Chor
chor@post.harvard.edu

This version: May 2001

Abstract

This paper constructs a data set on purchasing-power-parity (PPP) adjusted skilled and unskilled wages in 139 countries for the period 1968-1998, based on the International Labor Organization's (ILO) annual *October Inquiry* and the Freeman and Oostendorp (2000) *Occupational Wages Around the World* (OWW) file. It finds strong evidence for the existence of well-integrated markets for skilled and unskilled labor, justifying the approach of constructing a skilled wage series and an unskilled wage series. Several significant results emerged from an analysis of a representative subset of 67 countries which provided unbroken coverage for 1970-1994: (i) there is striking evidence of unconditional β convergence in the skilled-unskilled wage ratio worldwide; (ii) this relative wage convergence was especially strong within a "club" of open economies, suggesting that Heckscher-Ohlin-Sameulson mechanisms might be at work; and (iii) there is a relatively weak pattern of σ convergence in unskilled real wages, implying that the claim of "Divergence, Big Time" (Pritchett 1997) has to be qualified when factor markets are studied instead of aggregate incomes.

I wish to thank Richard Freeman, Michael Murray, Remco Oostendorp, and especially Jeffrey Williamson for their helpful comments and suggestions. Freeman and Oostendorp provided generous access to their OWW data set, while Aart Kraay provided an extended version of the Penn World Tables real GDP per capita estimates. This work was originally submitted as part of a senior thesis at Harvard University.

1 Introduction

This paper addresses the ongoing debate on cross-country convergence from the perspective of labor markets. It constructs and analyzes a new data set on purchasing-power-parity (PPP) adjusted wages for skilled and unskilled labor in a large sample of countries from 1968-1998. Up till now, research on growth and convergence has concentrated almost exclusively on the behavior of aggregate variables such as GDP per capita or per worker; Baumol (1986), Abramovitz (1986), Barro and Sala-i-Martin (1995), Caselli, Esquivel and Lefort (1996), and others have established many important results on the long run evolution of aggregate incomes. This paper, however, shares the basic contention of Williamson (1995) and O'Rourke and Williamson (1999) that factor incomes ought to be studied for the additional leverage that they give to our understanding of convergence processes, illustrated so well by studies on Heckscher-Ohlin-Samuelson (HOS) effects. There are at least three good reasons for studying factor incomes.[1]

First, economic forces leave their mark on factor markets directly. Convergence or divergence between countries should therefore be more readily manifested and identifiable in factor price trends than in aggregate income statistics. This is especially true if the economic forces have an unbalanced or assymetric impact on returns to labor, skills, land, and capital. Romer (1994), Pritchett (1997), and others have found scant evidence of convergence in real GDP per capita on a global scale. Will the evidence from labor markets corroborate this result?

Second, real GDP per capita and per worker are coarse statistics. By averaging out all factor incomes (or equivalently, all sector value-added production) over the entire population or workforce, these aggregates tend to obscure wage movements that impact the welfare of individual wage earners. Around the mid-1970s, the United States began to experience a secular rise in earnings inequality, characterized by a bottoming out of unskilled wages (Katz and Murphy 1992; Freeman and Katz 1995). In the context of wages, the skilled-unskilled wage ratio increased dramatically. A cross-country study of this relative wage ratio can enhance our understanding of how convergence forces in a globalizing world have influenced these trends in wage inequality.

Third, movements in factor prices provide valuable clues for sorting out the sources of

[1] The following draws on Williamson (1995) and O'Rourke and Williamson (1999).

convergence. The different hypotheses offered to explain the evolution of country incomes often make distinct predictions on the returns to skilled and unskilled labor. If broad-based, non-sector-biased technological change were the main force driving growth, both skilled and unskilled wages would rise in tandem, and the skilled-unskilled wage ratio should remain fairly stable. In contrast, if open economy forces were at work, HOS thinking suggests that production and exports in rich, skilled labor-abundant countries should shift towards skill-intensive industries, thus raising the relative skilled wage; the exact opposite would be true for poor, unskilled labor-abundant countries. The skilled-unskilled wage ratio would move in opposite directions in the two types of economies, generating a convergence in the relative wage structure across borders. An analysis of factor prices would allow us to take a first step towards testing these competing hypotheses.

The above discussion highlights two distinct concepts of convergence – absolute wage convergence and relative wage convergence.[2] Absolute convergence is closely related to growth theory and the "convergence hypothesis" – the claim that there exist "powerful and persistent" (Baumol 1986, p. 1084) catch-up forces that narrow income differences across countries. Relative wage convergence is more closely related to trade and HOS theory. These two concepts are not equivalent: Absolute wages could diverge if technological progress took place at different rates across countries; but relative wages could still converge if HOS convergence forces were operating in a globalizing world.

This paper aims to fill the gap in the literature on wage convergence over the past 30 years by exploiting a new data set on skilled and unskilled wages. Section 2 briefly reviews the existing work on absolute and relative factor price convergence, as well as two pioneering studies (Davis 1992; Freeman and Oostendorp 2000) which have attempted wage comparisons across a broad sample of countries. Section 3 documents the construction of the data set. The raw data is from the International Labor Organization's (ILO) annual *October Inquiry*, a survey of wages in 159 different occupations, which was recently calibrated for cross-country consistency in Freeman and Oostendorp's (2000) *Occupational Wages around the World* (OWW) file.[3] Table 1 lists the 139 economies in the data set and the respective years of coverage. The geographical scope offered is clearly much larger than for most other

[2]Throughout this paper, unless otherwise stated, the relative wage refers to the ratio of skilled to unskilled wages. This skill premium serves as a measure of wage inequality.

[3]I am much indebted to Richard Freeman and Remco Oostendorp for allowing the generous use of an extended version of the OWW for 1968-1998 before it was made available for public use. The manual entry of raw data from the *October Inquiry* for the years 1968-1983 was done by myself in November 1999, and Oostendorp calibrated a preliminary version of the extended OWW in January 2000.

existing data sets. Section 4 presents several results. Significantly, we find unconditional β convergence in the skilled-unskilled wage ratio. This relative wage convergence was even more impressive within a subset of 34 countries classified as open by Sachs and Warner (1995), suggesting that HOS forces may be at work in influencing relative wages around the world. We also find some evidence of absolute σ convergence in unskilled real wages, a fact which the macro literature would never have uncovered. Section 5 concludes.

2 Review

2.1 Absolute Income Convergence

The "convergence hypothesis" has its roots in Gerschenkron's (1952) idea of the "advantages of backwardness", which posited that less developed countries (LDCs) have the potential to adopt the best-practice technologies of industrial leader nations. As this improves worker productivity in LDCs, real wages should join real GDP per worker in converging on developed nation levels. Although this convergence should be more readily identifiable in factor price trends, the empirical work has focused overwhelmingly instead on the behavior of aggregate income variables, partly because of the availability of macro data sets such as Maddison (1995) and the Penn World Tables. It is now well-known that there has been little absolute convergence on a global scale in real GDP per capita or per worker (De Long 1988; Romer 1994; Pritchett 1997; Prados 2000; Bourguignon and Morrisson 2000). In fact, the OECD expanded moderately while parts of the Third World suffered negative growth, leading Pritchett (1997) to conclude that "Divergence, Big Time" has been the "dominant feature of modern economic history" (p. 3).

Instead, convergence is now viewed as a conditioned phenomenon. Incomes have converged within smaller "clubs" of countries, such as the OECD (Baumol 1986; Abramovitz 1986) or the open economies (Sachs and Warner 1995). In addition, regressions explaining growth rates as a function of the initial income and a host of conditioning variables (such as investment, education, openness, and the quality of government) have consistently yielded a negative coefficient on initial income, confirming the presence of conditional convergence once other country characteristics have been controlled for.[4] The conditioning variables can be viewed either as determinants of the country-specific steady state income level in

[4] Durlauf and Quah (1998) provides an excellent summary of such work. See also Barro and Sala-i-Martin (1995) and La Porta et al. (1999).

the Solow (1956) growth model,[5] or as elements of the "social capability" of an economy to exploit advanced technologies (Abramovitz 1986).[6]

In contrast to the extensive literature on aggregate incomes, work on the absolute convergence of wages has been limited. Long run wage convergence is best documented for the Atlantic Economy (comprising much of the present-day OECD). During the First Globalization Boom of 1870-1913, the coefficient of variation (CV)[7] of unskilled wages for a sample of 17 Atlantic economies fell dramatically by a third, due mostly to the shrinking wage gap between the New and Old Worlds. As transport costs declined and as labor migrated *en masse* from the Old World to the labor-scarce New World, both commodity and factor prices moved towards parity. Significantly, these trends were reversed during the retreat to autarky of the Inter-War years (1913-1950), illustrating the pivotal role of globalization in facilitating convergence (Bordo et al. 1999; O'Rourke and Williamson 1999). Beyond this, however, the evidence on wage convergence is sparse. We do know that the CV of unskilled wages in the Atlantic Economy (plus Argentina and Brazil) declined slightly after World War II (Williamson 1995), and we also have some data on unskilled wages in several Asian, Latin American, and Mediterranean Basin countries pre-1940 (Williamson 1999). But our understanding of convergence is incomplete due to the lack of data particularly for LDCs post-1950. For skilled wages, the evidence is even more piecemeal. Abowd and Boganno (1995) construct some data on executive and managerial compensation in the OECD, but we again have very little information from LDCs. Clearly, much work remains to be done in documenting the post-War global evolution of wages.

2.2 Relative Factor Price Convergence

Relative factor price convergence is a distinct concept from absolute convergence, in that the former is more closely linked to trade and HOS theory. HOS reasoning suggests that the skilled-unskilled wage ratio should rise in countries that are better endowed with skilled labor (relative to unskilled labor), as these countries exploit their comparative advantage in

[5]The Solow growth model predicts that aggregate income per capita, Y, converges to a country-specific steady state level. This extends readily to a prediction of convergence in wages to a steady state level. Denote skilled and unskilled labor by H and L respectively. If we assume that the marginal products $\frac{\partial Y}{\partial H}$ and $\frac{\partial Y}{\partial L}$ are continuous, then these will also converge to some steady state levels.

[6]An alternative response to the lack of convergence has been to reject the convergence hypothesis. One of the motivations for endogenous growth theory has been to construct models that can generate divergence in the distribution of cross-country incomes (Romer 1986; Lucas 1988; Aghion and Howitt 1992).

[7]The CV is the standard deviation divided by the mean to control for scale effects.

the production and export of skill-intensive goods. The opposite should occur in unskilled labor-abundant LDCs: The relative wage should decrease as production shifts to make use of unskilled labor more intensively. Together, this pair of forces generates a convergence in the relative wage across borders. In fact, if we further assume that countries share identical homothetic preferences and production technologies, HOS theory predicts a strict outcome of relative factor prize equalization.[8]

While HOS theory has motivated many studies on the impact of globalization on within-country wage inequality, much of this research has focused on the labor market structures in individual countries, and few studies have taken on a broader cross-country perspective. The case that has been studied most is the United States. From 1979-1989, the ratio of hourly earnings of full-time American workers in the 90th relative to the 10th percentile of the earnings distribution rose by 20% for males and 25% for females, consistent with the HOS predictions for skilled-labor abundant countries. Moreover, the erosion of the wage position was both relative and absolute, as the hourly earnings of young men with less than 12 years of schooling fell by 20% (Katz and Murphy 1992; Freeman and Katz 1995). The extent to which globalization has been responsible for this widening wage gap continues to be disputed. Some estimates have attributed as much as 50% of the rise in inequality in America to trade-induced effects (Wood 1994);[9] others have suggested more conservative estimates of 10-15% (Freeman 1995) or 15-33% (Feenstra and Hanson 1999), instead placing more credence on a skill-biased technological change explanation as the main force driving up wage gaps (Lawrence and Slaughter 1993; Berman, Bound and Griliches 1994; Cline 1997). Moving beyond the United States, there is some evidence on wage inequality in several other OECD countries from the collection of papers in Freeman and Katz (1995). In addition, the rise in wage inequality in Latin America in the 1970s has been contrasted with the fall in wage gaps within Asian LDCs such as Korea, Singapore and Taiwan (Wood 1997; Wood 1998). Nonetheless, few studies have to date sought to test explicitly for relative factor price convergence across a broad sample of countries, even though this is a key prediction which must hold if HOS effects have indeed played a role in influencing wages.

[8]See Dixit and Norman (1980) for a proof of this result.

[9]Wood (1994) estimated that the factor content embodied in net imports caused a 20% rise in the demand for skilled relative to unskilled labor. Given that the actual increase in demand was about 40%, this implied that 50% of the rise in wage inequality was caused by trade-related effects (Freeman 1995).

2.3 A Brief Survey of Cross-Country Wage Studies

Few comparative wage studies with a truly global scope have been published for the post-World War II period, primarily because of the lack of data from LDCs. Two studies stand out as exceptions. Davis (1992) was perhaps the first to include middle-income countries (Brazil, Colombia, South Korea, and Venezuela). Between 1960-1990, Davis documented rising earnings inequality in 7 advanced countries, with the converse trend of falling inequality in the 4 developing economies, a result consistent with HOS theory. However, Davis relied on population surveys, most of which were conducted by domestic statistical agencies with survey procedures differing from country to country, making it difficult to make precise comparisons of the figures across countries. Some country data sets reported earnings (inclusive of benefits and allowances), while others reported baseline wages. In some countries, Davis had to merge surveys conducted by different agencies to piece together a time trend. The use of micro data is thus only useful in the case of high or middle income countries with fairly uniform statistical practices.

Instead, this paper uses the occupational wage series from the OWW file (Freeman and Oostendorp 2000). Freeman and Oostendorp presented some evidence on the large dispersion (and hence the lack of concergence) in the wages paid for the same occupations in different countries. The authors also constructed a measure of the skill differential in each country (a measure of the observed within-country wage dispersion), and found an inverse U-shape Kuznets-type relationship between the skill differential amd GDP per capita. While Freeman and Oostendorp have made a major empirical contribution, the research potential of the OWW has yet to be exhausted. For a start, the longer panel from 1968-1998 used in this paper will allow us to tackle questions related to long run convergence. Also, Freeman and Oostendorp analyzed the OWW from the perspective of labor economists, thus paying special attention to the dispersion of wages across the full range of reported occupations. This paper focuses instead on growth and trade issues, and is thus disposed to thinking in terms of two broad forms of labor inputs, namely skilled and unskilled, and in terms of stylized wage measures such as the skilled-unskilled wage ratio. Finally, the wage inequality and skill differential measures used by Freeman and Oostendorp are, strictly speaking, not perfectly comparable across countries and time, because the set of occupations for which wages were actually reported varies substantially from country to country and from year to year. This paper shall instead use trends in skilled and unskilled wages from a relatively small, but consistent, subset of occupations from the OWW.

3 Deriving the Real Wage Data

3.1 Background

Our primary data source is the ILO's *October Inquiry*. Each year, the ILO solicits information on occupational wages and normal hours of work from country statistical agencies, and publishes these in a *Bulletin of Labor Statistics* supplement. Data for the month of October is requested, helping to control for possible seasonal effects.[10] Two factors make the *Inquiry* especially useful for comparative wage studies. First, the range of occupations and countries is very extensive. From 1953-1983, 48 occupations were surveyed; this was expanded in 1983 to a total of 159 occupations, drawn from the full spectrum of agriculture, manufacturing, and service industries. Up to 1998, an impressive total of 182 countries and territories had been represented in at least one issue of the *Inquiry* (Table 1, Freeman and Oostendorp 2000). Second, each occupation is clearly defined by a specific set of job tasks, as enacted in the ILO's *International Standard Classification of Occupations* (ISCO, 1968). This should in principle ensure that the country agencies submit wage returns for identical work, bringing us close to the ideal of consistent reportage across all countries.

However, the *October Inquiry* in its original form has one significant shortcoming: Different countries report wage levels using different earnings concepts, making it hard to make careful cross-country comparisons. For example, in 1995, the United States reported median weekly earnings; Germany listed collectively bargained wage rates; while India reported minimum wage rates (Freeman and Oostendorp 2000). In practice, "wage rates" refer to returns per unit time of labor effort, while "earnings" typically include additional compensation such as allowances or bonuses. While most of the figures were for both men and women, some countries reported wages for only one gender.[11] Freeman and Oostendorp (2000) dealt with this inconsistency problem by calibrating the raw *Inquiry* data econometrically in the OWW file, re-expressing all the figures into one standardized pay measure,

[10] The choice of October over other months appears to be purely accidental. A complete list of reporting agencies is in the appendices of the *Inquiry* for each post-1983 year.

[11] A quick note on alternative sources of wage data is in order. One alternative is the United Nations Industrial Development Organization (UNIDO) files, which report remuneration in manufacturing for a sample of countries similar to the *Inquiry*. The Bureau of Labor Statistics has also published a data release on "hourly compensation costs for production workers in manufacturing" for newly industrializing and developed countries, available at `http://stats.bls.gov/flsdata.htm`. These two data sets are used in Rodrik (1999). Unfortunately, both of these alternatives focus only on relatively unskilled labor. The UNIDO files are also plagued by inconsistencies, since the pay concepts reported there also range between wages and earnings.

the average hourly wage rate of male workers. The authors worked with a specification in which the log wage is a function of a matrix of pay measure, gender, occupation, and time dummies, plus country by year interaction terms, from which the average male hourly wage rate could be predicted. This was estimated using generalized least squares (GLS) to allow for a more complex error structure, as there are potentially different sources of error arising from variation in the different sets of dummy variables.[12]

3.2 Choice of Occupations

We chose 7 occupations each for skilled and unskilled labor from the OWW as the basis for constructing our real wage series. The approach of studying market wage trends by bringing together data from various industries has a long precedent in the labor literature (Slichter 1950). The 7 unskilled occupations selected were: thread and yarn spinners in the textiles industry (#25); sewing machine operators in the manufacture of wearing apparel excluding footwear (#30); laborers in printing, publishing and allied industries (#51); laborers in the manufacture of industrial chemicals and other chemical products (#56/#59)[13]; laborers in the manufacture of machinery except electrical (#70); laborers in electric light and power (#80); and laborers in construction (#90).[14] These choices satisfied three criteria. First, the job scopes did not require more than primary education. Second, the industries picked were found in most economies, ensuring wide geographical coverage. Third, to examine the impact of openness on wages, several of these jobs were in industries that manufacture potentially tradable goods, such as textiles and apparel. These 7 occupations lie on the low end of the wage spectrum in the OWW: In countries that listed wages for at least 80 of the 159 occupations during 1983-1998, the 7 occupations were in the lower one-third of the distribution of reported wages in at least 75% of country-year pairs, with one exception (#80).[15]

[12] Specifically, Freeman and Oostendorp (2000) estimated the following: $W_{i,j,o,t} = D_{i,j,o,t}A_{i,j,o,t} + \theta_{i,j,o,t}$, where i denotes the pay concept, j denotes gender, o denotes occupation, t denotes time, and $\theta_{i,j,o,t}$ is a random error term. $W_{i,j,o,t}$ denotes the log wage, $D_{i,j,o,t}$ is a matrix of pay measure, gender, occupation, time dummies, and country by year interaction terms, and $A_{i,j,o,t}$ is a vector of coefficients.

[13] Prior to 1983, the *Inquiry* reported only one wage figure for laborers in the chemical industry. This was broken into two occupations in 1983, for the manufacture of industrial chemicals (#56) and the manufacture of other chemical products (#59). The OWW links the series before 1983 with an unweighted average of #56 and #59.

[14] Parentheses contain the code number assigned to each occupation in the *October Inquiry*.

[15] Occupation #80 was in the lower one-third of the wage spectrum in 58% of country-year pairs. When countries do not report wages for all 159 occupations, it may be difficult to assess the relative position of the occupations on the full skill hierarchy, as the sample reported might have given more coverage to say the unskilled occupations. We try to minimize this effect by restricting the statistic to countries which

For skilled labor, the 7 occupations were: chemical engineers in the manufacture of industrial chemicals (#52); power distribution and transmission engineers (#76); bank accountants (#129); computer programmers in the insurance industry (#133); government executive officials in public administration (#139); mathematics teachers at the third (tertiary) level (#145); and general physicians (#152). The skilled workers we focus on are professionals, as opposed to artisans or craftsmen, who have been more closely documented in the economic history literature.[16] The "skilled" wages are thus a wage return to technical expertise that would require at least a secondary level of schooling. Certainly, these 7 occupations lie above the 75th percentile of the wage distribution for country-year pairs reporting at least 80 occupations during 1983-1998. Unfortunately, skilled occupations like these 7 were only introduced into the *October Inquiry* in 1983; we circumvent this lack of coverage in Section 3.4 by using proxies.

For the 14 occupations, nominal wages were deflated by a PPP index for consumption goods, and re-based in 1990 US dollars, the numeraire for this paper. A PPP index for consumption goods is used, since this most closely reflects the actual purchasing power over goods that workers themselves buy.[17]

Table 2a reports some summary statistics: the 5th, 50th, and 95th percentiles for the unskilled occupations, pooled over time. The distribution of wages in these 7 jobs were very similar, with only slight differences in the degree of right skew. The same observations apply also to the 7 skilled occupations (Table 2b). The average pair-wise correlation is 0.96 for the 7 unskilled jobs, and 0.93 for the 7 skilled jobs. Clearly, trends among the unskilled (respectively skilled) occupations are very similar, implying that there are well-integrated factor markets for unskilled and skilled labor. This justifies the approach of modeling the two forms of labor as separate factors of production with their distinct wage trends.

reported wages for at least 80 of the 159 occupations. We keep #80 because of the uniformity in job description for unskilled laborers in the ISCO (1968).

[16]The choice of #139 is defensible as government executives should earn a wage commensurate with market rates in the private sector to cover the opportunity cost of working in the public sector. Table 2b confirms that the distribution of wages for #139 was similar to that of the other skilled occupations.

[17]The PPP index reports the amount of local currency equivalent in purchasing power over consumption goods to one 1990 US$. PPP figures up to 1992 were from the Penn World Tables. These were extended to 1998 using the formula $PPP_{t+1} = PPP_t \times \frac{CPI_{t+1}}{CPI_t}$, where PPP_t and CPI_t denote respectively the PPP index and the consumer price index in year t. CPI data was from the "Global Development Network Growth Database" (Easterly and Yu 1999). Where necessary, this was augmented using the IMF's *International Financial Statistics Yearbook* (IFS). These two sources are almost completely identical for overlap years.

3.3 An Econometric Framework

We now outline the econometric framework for constructing the unskilled wage series. (The entire exposition carries over to the skilled wage series as well.) Let $w^U_{c,t}$ denote the underlying wage series for the base-line market returns to unskilled labor, free of occupation-specific effects. Here, c denotes country, and t denotes year. Let $w^i_{c,t}$ denote the observed real wages in the 7 unskilled occupations, where i ranges from 1 to 7. We posit that each observed real wage series is a linear function of the underlying wage trend $w^U_{c,t}$:

$$w^i_{c,t} = \alpha^i_1 + \alpha^i_2 w^U_{c,t} + \varepsilon^i_{c,t} \tag{1}$$

Notice that α^i_1 and α^i_2 are specific to each occupation, while $\varepsilon^i_{c,t} \sim N(0, \sigma^2_{\varepsilon_i})$ is independent normally distributed random error. The system in (1) is seriously under-identified, with twice as many coefficients as there are equations. While we would like to estimate $w^U_{c,t}$ directly with $E(w^U_{c,t}) = -\frac{\alpha^i_1}{\alpha^i_2} + \frac{1}{\alpha^i_2} w^i_{c,t}$, this is not possible given that estimates of all the α's cannot be obtained. Instead, we run pair-wise regressions of one occupation against another, to tease out the trends in occupation i that can be predicted by movements in occupation $j \neq i$. In this way, we capture the commonalities in wage trends that are the underlying factor price trends in the unskilled labor market.[18] In particular, we solve (1) simultaneously for occupations i and j to eliminate $w^U_{c,t}$:

$$w^i_{c,t} = \left(\alpha^i_1 - \frac{\alpha^i_2 \alpha^j_1}{\alpha^j_2}\right) + \frac{\alpha^i_2}{\alpha^j_2} w^j_{c,t} + \left(\varepsilon^i_{c,t} - \frac{\alpha^i_2}{\alpha^j_2}\varepsilon^j_{c,t}\right), \qquad 1 \leq i, j \leq n \tag{2}$$

It is crucial that the correlation between $w^i_{c,t}$ and $w^j_{c,t}$ should derive from the common wage trend, $w^U_{c,t}$, rather than from any correlation between the error terms, $\varepsilon^i_{c,t}$ and $\varepsilon^j_{c,t}$. If $Cov(\varepsilon^i_{c,t}, \varepsilon^j_{c,t}) \neq 0$, the apparent correlation between $w^i_{c,t}$ and $w^j_{c,t}$ would potentially be spurious, invalidating the claim that there exists a common wage trend. The most likely sources of correlation between the error terms are country effects, such as localized economic events or biases in the reporting practices of domestic statistical agencies. To control for these, a random effects specification was considered, in which the error term in (2) was re-written as: $\left(\varepsilon^i_{c,t} - \frac{\alpha^i_2}{\alpha^j_2}\varepsilon^j_{c,t}\right) = \mu^i_c + \nu^i_{c,t}$, where $\mu^i_c \sim N(0, \sigma^2_\mu)$ is a country-specific error term, and $\nu^i_{c,t} \sim N(0, \sigma^2_\nu)$ is pure random error. However, when the Hausman test was performed, it overwhelmingly favored a fixed effects model: The coefficient of $w^j_{c,t}$ from fixed effects estimation was significantly different at the 5% level from that obtained from random effects estimation (feasible GLS) in 37 of the $7 \times 6 = 42$ pair-wise regressions among

[18] I am greatly indebted to Michael Murray for this suggestion.

unskilled occupations. The same was found for 39 of the 42 skilled wage regressions. We thus re-write (2) with country fixed effects:

$$w^i_{c,t} = D_c \beta^{i,j}_c + \frac{\alpha^i_2}{\alpha^j_2} w^j_{c,t} + \left(\varepsilon^i_{c,t} - \frac{\alpha^i_2}{\alpha^j_2} \varepsilon^j_{c,t} \right) \quad (3)$$

Here, D_c is a matrix of country dummies which replaces the constant term in (2), and $\beta^{i,j}_c$ is a vector of fixed effects coefficients.[19]

The regression results for (3), run using the interpolated occupational wage series, are reported in Tables 3a and 3b. All R^2 values, with the odd exception, are high, averaging 0.86 and 0.55 respectively for the unskilled and skilled occupations.[20] Most of the variance of $w^i_{c,t}$ in (3) is thus accounted for by the dependent wage series, $w^j_{c,t}$, and by the country dummies; the $\varepsilon^i_{c,t}$ and $\varepsilon^j_{c,t}$ terms are small, insofar as they do not significantly obscure the common trends that underlie the occupational wage series. This once again confirms that markets are well-integrated for both skilled and unskilled labor.

Using $D_c \beta^{i,j}_c + \frac{\alpha^i_2}{\alpha^j_2} w^j_{c,t}$ as a linear predictor, we obtain 7 series that capture trends in $w^i_{c,t}$ – the original series and 6 series predicted from regressions. For countries which did not report $w^i_{c,t}$ but did report $w^j_{c,t}$, we are unable to compute this linear predictor since the country coefficients could not be estimated. In such cases, out-of-sample country effects were estimated from a regression of the country coefficients in $\beta^{i,j}_c$ against the mean predicted value of $\frac{\alpha^i_2}{\alpha^j_2} w^j_{c,t}$. This uses the fact that in a fixed effects model, the country coefficients are correlated with the mean predicted wage net of the fixed effects. This procedure may be *ad hoc*, but any errors introduced are small, as indicated by the average standard error of the predictors (SEP) for the out-of-sample effects (reported as the first "SEP" figure in the cells in Tables 3a and 3b.) In Table 3a, the average of these 42 SEP figures was 0.10, barely 3.9% of the average standard deviation of US$2.66 for wages in the 7 unskilled occupations. Similarly in Table 3b, this average SEP figure was 0.78, or 10.4% of the average standard deviation of US$7.49 for wages in the 7 skilled occupations.

For each i, the 6 series predicted from (3) were highly correlated with the original $w^i_{c,t}$, with the correlation coefficient always larger than 0.95 for unskilled wages and 0.94 for skilled wages. We took a simple unweighted average of these 7 unbiased estimators of $w^i_{c,t}$

[19]The possibility of time effects was considered, namely that common economic shocks across countries might induce systematic shifts in wage levels in a given year. However, time dummies were rarely significant in fixed effects regressions, and did not add much to the R^2's. The time dummies were therefore dropped.

[20]The R^2's for the skilled job regressions was smaller due to the presence of more outliers.

to obtain a series $\hat{w}^i_{c,t}$ that incorporates unskilled wage trends from all 7 occupations. The average pair-wise correlation for the 7 $\hat{w}^i_{c,t}$'s was 0.95 for unskilled occupations and 0.78 for skilled occupations, indicative of the common factor price trend in each labor market. To back out a single unskilled wage series, one would ideally take a weighted average of the $\hat{w}^i_{c,t}$'s, using the employment shares of each occupation i as weights. As such figures were not readily available, we settled for a simple unweighted average instead. While this unweighted average may be shifted from the true level of $w^U_{c,t}$ by a constant, it nonetheless captures the salient wage movements which are the focus of our attention.[21]

One last econometric complication remains to be addressed. In (3), the covariate $w^j_{c,t}$ is measured imprecisely and this introduces issues pertaining to measurement error. Since $Cov\left(\frac{\alpha^i_2}{\alpha^j_2}w^j_{c,t}\ ,\ \varepsilon^i_{c,t} - \frac{\alpha^i_2}{\alpha^j_2}\varepsilon^j_{c,t}\right) = -\left(\frac{\alpha^i_2}{\alpha^j_2}\right)^2 Var(\varepsilon^j_{c,t}) \neq 0$, the assumption of least squares models requiring independence between the error terms and the covariates is violated. The estimated coefficient of $w^j_{c,t}$ is thus asymptotically inconsistent, with an attenuation bias towards zero.[22] For our regressions, however, this bias is not serious. The standard error of the linear predictors, $D_c\beta^{i,j}_c + \frac{\alpha^i_2}{\alpha^j_2}w^j_{c,t}$, are all tolerably small (reported as the second "SEP" figure in each cell in Tables 3a and 3b.) The average SEP was 0.19 in Table 3a and 1.04 in Table 3b, respectively 7.1% and 13.9% of the average standard deviation of wages in the 7 unskilled and skilled occupations. The imprecision introduced by the residual errors is thus relatively small, though slightly larger for the skilled regressions. Moreover, all the slope coefficients in Tables 3a and 3b are statistically significant at the 5% level, with only 2 exceptions.[23]

3.4 Proxying for the Full Skilled Wage Series

The work in section 3.3 has produced an unskilled wage series for 1968-1998, and a skilled wage series for 1983-1998. To extend the skilled wage series back to 1968, 10 "fairly skilled" occupations for which coverage in the *October Inquiry* spanned all 30 years were used as proxies. Table 2c lists these occupations, along with summary statistics. By comparing the median with that from Tables 2a and 2b, one can see that these 10 jobs lie

[21]Our constructed unskilled wage series is consistent with the O'Rourke and Williamson (1999) Atlantic economy data base. For overlapping data points, the correlation between the two series is 0.61.

[22]See Greene (1997, p. 439-440) for a discussion of this problem.

[23]The exceptions are the pairs #52 & #129, and #52 & #152. Measurement error should be more serious for the skilled wage series because of the larger number of outliers. However, the attenuation bias should reduce the point estimates for the skilled wages by more than for the unskilled wages, which would lead the skilled-unskilled wage ratio to under-state wage inequality.

between the skilled and unskilled occupations in the wage hierarchy. These fairly skilled occupations require some literary or technical proficiency, though they fall short of being fully professional jobs. Within the predominantly unskilled set of occupations surveyed pre-1983, for countries reporting at least 24 of the 48 wages requested, the fairly skilled occupations were higher than the median wage in at least 60% of the country-year cells, with two exceptions (#91 and #92).[24]

The estimation proceeds much as before, with the skilled wage series for 1983-1998 now regressed on the fairly skilled occupations one at a time, together with country fixed effects.[25] Out-of-sample country effects are estimated using the same procedure in Section 3.3. Summary results are presented in Table 3c. Importantly, the R^2's obtained are all high, ranging from 0.59 to 0.79; our proxies therefore account for a large share of the variance in skilled wages, larger perhaps than one might initially have expected given that there are non-trivial discrepancies in skill intensities between the two sets of occupations. Furthermore, the slope coefficients are all significant at the 1% level. Once more, the imprecision due to possible measurement error is reasonably small: The average of the 10 SEPs for the linear wage predictor is 0.49, or 14.8% of the average standard deviation of US$3.33 for the reported wages in the 10 proxy occupations. (The average of the 10 SEPs for the out-of-sample coefficient predictors is an even smaller 0.45.) The coefficients estimated from the 1983-1998 data are then used to obtain 10 predicted skilled wage series for the full period. This necessarily assumes that the relationship between the fairly skilled and skilled occupations was stable throughout 1968-1998. Finally, we take a precision-weighted average of the 10 predicted series to obtain the full skilled wage series; here, the precision is the reciprocal of the variance of each linear predictor.[26]

Reassuringly, there is a tight fit between the proxied wage series and the skilled wage series constructed in Section 3.3 for the 821 data points from the overlap years 1983-1998. The correlation between the two series is 0.99. The mean square error between these two series is a relatively small US$0.70, a figure that is inflated primarily because of a handful of outliers. The median square error is an even smaller US$0.06. The rest of this paper uses the fully proxied skilled wage series for 1968-1998.

[24] #91 and #92 were above the median in 56% and 34% of the country-year cells respectively.

[25] The Hausman test comes up significant even at the 1% level for 3 of these proxies. For consistency, a fixed effects model as in (3) is used for all 10 regressions.

[26] Given independent unbiased estimators $\hat{\theta}_i$ for a quantity, the minimum variance unbiased estimated that is also a linear combination of the $\hat{\theta}_i$'s is the precision-weighted average. This result follows from solving a standard constrained maximazation problem via a Lagrangian.

4 Analysis of Wage Trends (1970-1994)

We now analyze the real wage series constructed in Section 3, focusing on the sample of 67 countries asterisked in Table 1. This is the largest subset of the 139 countries for which unbroken wage series for both skilled and unskilled labor could be derived for 1970-1994 with a minimal amount of extrapolation.[27] Here, we are trading off consistency for scope, since wage levels in this subset can be more easily compared over time without having to worry about countries entering or dropping out of the sample. Nonetheless, we do not lose much in terms of the generality of the results. The smaller panel includes countries from all major geographical regions and levels of economic development, while the analysis repeated for the full sample of 139 economies yields almost identical results.[28]

4.1 The Worldwide Distribution of Real Wages

We start with an overview of the salient features of the distribution of real wages. Most significantly, there is a large amount of dispersion in wage levels across countries. In 1994, the highest skilled wage in the sample (Trinidad and Tobago) was 21 times the lowest (Sudan). The disparity was even greater for unskilled labor: The highest-paid unskilled workers (Sweden) earned 27 times more than the lowest (Zambia). We are clearly far from a world of strict equalization of factor prices. Figure 1 illustrates this point with kernel density plots of the distributions of skilled and unskilled wages in 1994.[29] Notice that both diagrams are right-skewed; as recently as the 1990s, workers in high-income countries continued to earn significantly more than their counterparts in LDCs. For skilled wages, the plot has a wide-based unimodal shape (Figure 1a). Kernel density plots from earlier years are essentially similar, indicating that there have been few dynamic changes in the shape of the skilled wage distribution since 1970.

More interestingly, for unskilled wages, there is some evidence for the emergence of a "twin-peaked" distribution, along the lines of Quah's (1996; 1997) hypothesis for the global distribution of real GDP per capita. Kernel density plots from earlier years such as 1970 and 1982 show a much less distinct bimodal shape, resembling more a unimodal shape with a long and uneven right tail. This evidence on "twin peaks" is suggestive of the emergence

[27] This involved geometrically extrapolating the data forward and backward in time by 3 to 4 years. A few countries for which the extrapolations suggested unreliably large growth rates were dropped.

[28] This parallel analysis is available upon request.

[29] The `kdensity` command in `STATA` was used to produce these plots.

of two convergence clubs, with economies within each club becoming increasingly similar, possibly because of increased interaction through trade or technological transfers. In Figure 1b, we observe one distinct peak centered near the US$2 hourly wage mark, comprising mainly Third World economies; a smaller peak around the US$6.50 level, consisting mostly of fast-growing LDCs; and a long right-tail made up of the OECD countries.

We briefly examine the evolution of average wage levels by geographical regions over the entire 25 year period.[30] Among less economically developed regions, wage growth was strongest in the Pacific Basin and in North Africa and the Middle East. Between 1970-1994, unskilled hourly wages rose from US$3.00 to US$4.22 in the Pacific Basin, and from US$1.72 to US$3.26 in North Africa and the Middle East. Conversely, unskilled wages in North America and the Caribbean fell from US$17.42 to US$16.00, contributing to an impressive convergence across the Pacific Ocean. Unskilled wages in the United States alone eroded more than 50% between the mid-1970s and the early 1990s.[31] Elsewhere, however, there was little change in the wage hierarchy from 1970-1994. The high wages in Western Europe continued to increase, registering per annum growth of 0.28% for skilled wages and 1.26% for unskilled wages.[32] In contrast, skilled and unskilled wages in South and Central America, South Asia, and Sub-Saharan Africa recorded little or negative growth, and remained stuck at the bottom of the pecking order.

There are three points worth highlighting on the relative skilled-unskilled wage. First, wage inequality increased dramatically in several Third World regions. In South and Central America, this wage ratio surged from 3.1 in 1970 to a high of 4.3 in 1985 following the Latin debt crisis. In Sub-Saharan Africa, the relative wage actually fell initially, but this trend reversed in the late 1970s when aggregate incomes also began to collapse in many African countries (Pritchett 1997), leaving it as the region with the highest level of wage inequality by 1994. Second, relative wages converged between the Pacific Basin and North America. The skilled-unskilled wage ratio fell more than 40% in the Pacific Basin between 1970-1994, from 5.1 to 3.0, to draw even with North American levels. Third, wage inequality in Western Europe was among the lowest in the world, and even declined slightly. In

[30] More details on wage trends and growth rates by geographical regions for both skilled and unskilled wages are available upon request. The classification of countries by regions follows Table 1.

[31] This 50% drop is admittedly much larger than corresponding figures in the labor literature, which place the decline in American unskilled wages in the order of 20% (Freeman and Katz 1995). However, the wage series in this paper captures only baseline wage rates, and omits overtime pay and other earnings that are usually recorded in the micro labor surveys.

[32] The average annual growth rate was computed using the standard formula: $\frac{1}{T}(\ln(w_T) - \ln(w_0))$, where T is the time elapsed in years, and w_T and w_0 are the final and initial wage levels respectively.

contrast, the relative wage ratio rose from 2.5 to 3.0 in the United States, a 20% rise in wage inequality consistent with findings in the labor literature. The American experience of rising inequality was thus not shared by all OECD countries, as previously noted by Katz, Love and Blanchflower (1995).

4.2 β Convergence

We formally assess the strength of convergence forces in the spirit of William Baumol's (1986) article. Recall that unconditional β convergence exists when the growth rate of per capita income is negatively related to the initial income level; in other words, countries that start with lower per capita incomes tend to grow faster and catch up with richer countries. We therefore run regressions of the average growth rate in the wage (or income) variables between 1970-1994 against the initial 1970 wage (or income) level.

First, we perform this exercise for real GDP per worker to establish a benchmark for comparison.[33] Table 4, Column (1) shows that there is no evidence of β convergence in this aggregate income variable. In fact, the point estimate of the coefficient on initial income (the β coefficient) is positive though statistically insignificant, suggesting weak divergence. This reproduces the well-known result of Romer (1994). When we repeat this analysis for the skilled wage series, the positive β coefficient in Table 4, Column (2) is even significant at the 10% level. A convergence story only begins to emerge when we examine unskilled labor. The slope in Figure 2a is now negative, although this β coefficient is still not statistically significant even at the 10% level (Table 4, Column (3)). The evidence for unconditional β convergence in unskilled wages is thus weak and tentative. One should note that the OECD countries, with initial wages lying above US$4, are most responsible for tilting the slope of the regression line towards a negative gradient, suggesting that these economies form a distinct convergence club (Baumol 1986).

The most striking result pertains to the relative wage ratio. In Figure 2b, the data points cluster to indicate a distinct negative association between initial levels and subsequent growth rates. The β coefficient is significantly negative even at the 1% level (Table 4, Column (4a)). When two outliers, Burundi (BDI) and Nigeria (NGA), are deleted, the β

[33] Real GDP per capita in constant international dollars were from the Penn World Tables, extended past 1992 by Aart Kraay in a data set procured through personal communication. To convert the data into per worker form, the United Nations five-yearly demographic estimates were used. We take the workforce to be the population between ages 15 and 64. Missing years in the UN data base were geometrically interpolated. The income figures were re-expressed in 1990 US$.

coefficient almost doubles, indicating an even faster rate of convergence (Table 4, Column (4b)). We thus have strong evidence of relative wage convergence within this wide sample of 67 countries. Note that this finding does not contradict the fact that there was little absolute convergence in skilled wages and at most weak convergence in unskilled wages. Absolute wage levels can diverge even while relative wages converge if technological progress is Hicks-neutral, so long as the rates of technological progress differ across countries. Intuitively, skilled and unskilled wages would rise in tandem within countries, but at different speeds across countries, making it possible for the relative wage to converge.

We can quantify the rate of convergence by calculating a convenient "convergence half-life" – the length of time that needs to elapse for a given income gap to be reduced by 50%. We take this initial income gap to be the 5th-95th percentile gap in the 1994 distribution of the relevant wage (or income) variable. The last row of Table 4 calculates this "convergence half-life" for all regressions in which the β coefficient is negative.[34] For our best specification in column (4b), this half-life was approximately 33 years, in the same order of magnitude as the half-life that growth economists have obtained for empirical tests of the Solow growth model that report convergence rates of 2-3%.[35]

The unconditional β convergence in the skilled-unskilled wage ratio is a fairly surprising result, especially in the light of the empirical work of Daniel Trefler (1995). Trefler calibrated a generalized version of the HOS model – the Heckscher-Ohlin-Vanek model – that admits for an arbitrary number of countries and factors of production, and found that the observed factor-content of trade was significantly lower than that predicted by the Vanek equations, generating the "mystery of the missing trade". Our finding of unconditional β convergence is thus important: It suggests that while one does not observe relative factor price equalization, the insights of HOS theory nonetheless remain relevant for explaining this underlying tendency towards relative convergence. Remarkably, this convergence applies to a broad sample that includes economies with varying degrees of openness to international trade. On the other hand, the lack of relative factor price equal-

[34] Suppose that $y_i^* = \alpha + \beta(y_{i0}) + \varepsilon_i$, where i indexes countries, y^* is the growth rate of the wage (or income) measure y, and the subscript of 0 denotes the initial time period. Consider two countries indexed by i and j. Taking differences of this first equation yields: $y_i^* - y_j^* = \beta(y_{i0} - y_{j0})$. Let z_t be the difference in y values between i and j at time t. Re-express this latter equation as: $\frac{1}{z_t}\frac{dz_t}{dt} = \beta z_0$. Solving this differential equation and setting $z_t = 0.5z_0$ gives the "half-life", $t_{\frac{1}{2}} = -\frac{\ln 2}{\beta z_0}$.

[35] For example, Mankiw et al. (1992), and Caselli et al. (1996). Note that the half-life in the Solow-type regressions is the number of years required to halve the gap between the current income level and the steady state level. This half-life is not directly comparable with $t_{\frac{1}{2}}$, even though the orders of magnitude are suggestive of a similar rate of convergence.

ization can be attributed to the fact that the assumptions needed to generate this result, such as identical technologies and identical, homothetic preferences, are unlikely to hold in the real world.

4.3 A Convergence Club of Open Economies

HOS theory suggests that the trend towards relative wage convergence should be even stronger for countries that are open to international economic forces. Do we have evidence then that the open economies form a convergence "club"? Sachs and Warner (1995) provides a convenient "openness" variable for classifying economies. This criterion deems a country to be closed if it had one or more of the following characteristics during 1980-89: non-tariff barriers covering 40% or more of trade; average tariff rates of 40% or more; a black market exchange rate depreciated by 20% or more relative to the official exchange rate during the 1970s or 1980s (indicating an over-valued, tightly controlled official exchange rate); a socialist economic system; or a state monopoly on any major exports.

Admittedly, there are several criticisms of the Sachs-Warner criterion. The cut-offs of 40% and 20% are purely normative gauges of levels of protection that might be considered prohibitive. Also, there are strange anomalies. India during the 1990s is considered open in spite of an industrial sector notoriously resistant to foreign participation. China, on the other hand, is classified as closed, despite its booming trade volumes. We nonetheless use the Sachs-Warner criterion for two reasons. First, as a matter of practicality, the available alternatives all come with their own problems. The ratio of exports or imports to GDP is indicative of openness, but trade volumes are endogenous to growth rates, making it difficult to deduce a causal relationship between openness and growth. Moreover, these ratios are small (less than 15%) for OECD economies such as the United States which we would consider to be open. Other common measures of openness, such as tariff and non-tariff barriers, have already been incorporated in the Sachs-Warner variable. Second, we can specifically test the assertion of Sachs and Warner (1995) that the economies they classify as open constitute a convergence club. In all, 34 countries were classified as open, while 28 were considered closed; 5 economies could not be classified because of inconclusive primary source evidence on trade regimes.[36]

[36] The 34 open economies are: Mauritius, Ghana, Tunisia, Cyprus, Jordan, Bolivia, Guyana, Venezuela, Costa Rica, El Salvador, Guatemala, Mexico, Indonesia, Malaysia, the Philippines, Singapore, and all the OECD countries. The fact that this includes many LDCs implies that "openness" is not simply a proxy variable for high income countries. The 5 unclassified are: Seychelles, Suriname, Puerto Rico, St Lucia,

Figure 3 and Table 5 report the results on β convergence for the 34 open economies. As in Section 4.2, the evidence is unimpressive for real GDP per worker and skilled wages: The β coefficients in Table 5, Columns (1) and (2) are insignificant even at the 10% level, while the point estimate for skilled wages continues to suggest divergence. However, convergence was impressive for the remaining two wage variables, even more so than for the larger sample of 67. In Figure 3a, the negative relationship between the growth rate of unskilled wages and the initial level is now statistically significant even at the 1% level, with a computed half-life of 29 years (Table 5, Column (3)). The linear fit is even tighter in Figure 3b, which had the highest adjusted R^2 among all the regressions of 0.33 (Table 5, Column (4)). In short, these 34 open economies exhibited an even stronger trend towards unconditional β convergence for both unskilled wages and relative wages, suggesting a link between globalization and convergence. Sachs and Warner were right about the existence of a convergence club of open economies.

The evidence on relative convergence is strengthened if we consider a breakdown of the open economies into the OECD and non-OECD countries. In the non-OECD economies, the average skilled-unskilled wage ratio fell almost 25%, from a high of 5.0 in 1985 to 3.8 in 1994, towards the lower levels in the OECD countries (which hovered around the 2.4 mark during this period). This is again very appealing from the perspective of HOS theory, pointing to the convergence of relative wages between developed and developing countries within the larger set of open economies. It also suggests that most of the relative convergence was driven by the declining skilled-unskilled wage in the non-OECD open economies, rather than by a rising wage ratio in the OECD. Once again, the rise in wage inequality in the United States seems to have been the exception rather than the rule for the OECD. Unfortunately, HOS theory does not provide a ready explanation for this assymetry in the movement towards relative factor price convergence.

What about the direct links between openness and growth rates? Sachs and Warner argued that open economies had unambiguously higher rates of income per capita growth since 1965 than closed economies (Figure 2, Sachs and Warner 1995). Using geographic variables as instruments for the trade volume of a country, Frankel and Romer (1996) also showed that trade has a large positive effect on growth rates. To shed some light on this issue, Table 6 computes income and wage growth rates to allow us to make some comparisons. Looking at the far-right column, the 28 closed economies generally performed

and Iceland. The remaining 28 countries are classified as closed. Please see the appendix to Sachs and Warner (1995) for more details.

badly: Both skilled and unskilled wages recorded negative average annual growth rates (-0.71% and -0.86% respectively), while wage inequality rose by a non-trivial +0.93%. In contrast, the first column shows that both wage and real GDP per worker growth rates were unambiguously higher, while increases in wage inequality were unambiguously smaller, among the open economies than in the closed ones. There were also important differences in the growth dynamics within the set of open economies. For the non-OECD open economies, average growth rates for both skilled and unskilled wages drew level with the OECD during 1983-1994: For example, skilled wage growth saw a major improvement from -0.43% for 1970-1982 to +0.08% in the later period, virtually equal to the +0.09% experienced in the OECD. One should note that 10 of the 18 non-OECD open economies achieved full trade liberalization only in the mid-to-late 1980s – Ghana in 1985, Costa Rica in 1986, and the Philippines in 1988 to cite a few examples (Sachs and Warner 1995).[37] That the opening of trade regimes coincided with the catch-up of wage growth rates provides further evidence that trade has been benign to workers in these developing economies. Admittedly, this relationship may not be a causal one; open economic policies and healthy wage growth may both have been driven by some unidentified third factor, such as prudent government. Nonetheless, the strong correlation between the two merits further investigation.

4.4 σ Convergence

The β convergence measure is a long run phenomena, requiring about 30 years to shrink the 5th-95th percentile gap in 1994 by half. A more immediate measure is σ convergence, defined simply as a decline in the actual dispersion of wages (or incomes). It is important to note that these two convergence concepts are not equivalent. β convergence is a necessary but insufficient condition for σ convergence, since year-to-year fluctuations or shocks can increase the short-term cross-sectional dispersion of wages.[38]

The evidence on σ convergence is generally less dramatic than that for β convergence. Significantly, the CV of unskilled wages in the panel of 67 countries did decrease slowly but steadily from 0.89 to 0.83. This σ convergence is not particularly impressive by historical standards: Recall that during 1870-1913, the CV of unskilled wages fell by one-third within

[37] These countries were classified as open although they were fully liberalized for only part of 1970-1994. One should note that this generous definition of an open economy biases us against detecting convergence: With a wider sample of "open" economies, we risk adding countries to a potential convergence club that might not belong there in the first place.

[38] See Barro and Sala-i-Martin (1995) for a proof.

the Atlantic Economy, driven mostly by the migration of up to 60 million people across the Atlantic Ocean (O'Rourke and Williamson 1999). The last 30 years have not seen labor mobility anywhere close to these levels, which may partially explain why σ convergence has been weaker post-World War II. This muted σ convergence confirms similar results reported in Williamson (1995) for the Atlantic Economy plus a handful of Latin American countries. In contrast, the CV of skilled wages increased through 1970-1994 from 0.56 to 0.63. Last, but not least, the CV for the relative wage fluctuated a lot more, rising from 0.65 in 1970, to a high of 0.73 in 1986, dipping to a low of 0.60 in 1991, and then leveling off at 0.66 in 1994. In short, our wage series point towards σ divergence for skilled wages, weak σ convergence for unskilled wages, and no conclusive trend for the skilled-unskilled wage ratio.

Was the moderate σ convergence for unskilled wages more an inter-regional or an intra-regional phenomenon? To address this, we decompose the total worldwide variance of the unskilled wages into shares due to within-group and between-group variation (using the regional groupings in Table 1 and a standard sum of squares decomposition). Initially, in the 1970s, the between-group and the within-group components were approximately equal. Around the early 1980s, however, the between-group variance began rising steadily, while the within-group variance started falling. Eventually, by 1994, the between-group and within-group shares accounted for about two-thirds and one-third respectively of the total variance in unskilled wages. In short, convergence took place primarily within regions; as unskilled wage levels drew closer within geographical locales, this had the effect of increasing the variance of wages across regions.

We can pinpoint even more specifically where most of the convergence took place – among the open economies, and especially among the open non-OECD economies. For this latter group of LDCs, the CV of unskilled wages fell substantially between 1970-1994 from 0.93 to 0.60, a more than 30% decline. For the relative wage ratio, the CV actually rose from 0.48 in 1970 to a peak of 0.78 in 1986, but then fell by more than 50% to close off the period at a low of 0.35. It helps to remember that many of the countries within this category became open to international trade only in the late 1980s. Thus, the relative wage convergence coincided with policy moves towards liberalization. Even the CV of this wage ratio for the OECD countries, which typically had been open since the 1960s, experienced noticeable convergence, falling from 0.35 to 0.25. Conversely, the closed economies did not share in this relative convergence – the CV of the wage ratio for these 21 countries fluctuated cyclically during the 25 year period, but underwent no distinct trend. Moreover, the CV between open and closed economies more than doubled from 0.08 to 0.22, indicating a divergence in relative wage levels between the two groups. This uniqueness of the convergence experience to the open economies once again invites the conclusion that HOS forces were at work in pushing relative wages towards parity among open economies.

5 Conclusion

This paper has constructed a new standardized real wage data base for skilled and unskilled labor the world around over the past 30 years. In Section 4, we brought this new data to bear on various key issues in convergence and HOS theory. We found some evidence that unskilled wages underwent modest σ convergence over the last 30 years, despite the well-publicized results on aggregate per capita income divergence in the macro literature. The most striking result though was the uncovering of a broad-based relative convergence in the skilled-unskilled ratio worldwide. This trend was even stronger within a "club" of 34 open economies, suggesting that HOS forces may be at work in influencing wage premia in a globalizing world.

There is however still a lot of scope for research. A natural next step would be to try to quantify the sources of relative wage convergence more precisely. Movements in relative factor prices can in general be attributed to three sources – changes in factor endowments, changes in relative prices (which reflect the effects of trade and openness), and changes in technology (Jones and Engerman 1996). A regression approach has already been used to decompose the sources of movements in the wage-rental ratio in the Atlantic Economy of the First Globalization Boom (O'Rourke and Williamson 1999). To do the same for our relative wage series, additional data on skilled and unskilled labor factor endowments, and the relative prices of skilled labor-intensive and unskilled labor-intensive goods would be needed. Most importantly, we would need separate measures of technology or productivity in skilled labor-intensive and unskilled labor-intensive industries. This decomposition exercise could potentially help to answer such questions as whether HOS effects or skill-biased technological change have been the main cause of the rise in inequality in the United States. Also, it could allow us to test for a link between wage catch-up and technological catch-up to verify the Gerschenkron (1952) story explicitly.

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Table 1
Country Coverage by Geographical Regions (1968-1998)

Years of coverage are in round parentheses. The first range applies to unskilled wages, while the second range applies to skilled wages. The 67 countries analyzed in Section 4, for which skilled and unskilled wage series could be deduced for 1970-1994, are indicated by ***.

Sub-Saharan Africa (42)

Angola [AGO] (82-97; 82-97)
Benin [BEN] (68-91; 68-91) ***
Botswana [BWA] (75-84; 75-84)
Burkina Faso [BFA] (68-91; 68-91) ***
Burundi [BDI] (68-92; 68-92) ***
Cameroon [CMR] (68-92; 68-92) ***
Cape Verde [CPV] (76-86; 76-86)
Central African Rep. [CAF] (70-97; 70-97) ***
Chad [TCD] (70-97; 70-97) ***
Comoros [COM] (78-92; 78-92)
Congo, Dem. Rep. [ZAR] (69-78; 69-83)
Congo, Rep. [COG] (68-79; 68-79)
Cote d'Ivoire [CIV] (68-97; 68-97) ***
Djibouti [DJI] (78-87; 78-87)
Ethiopia [ETH] (68-78; 68-94)
Gabon [GAB] (68-95; 68-95) ***
Ghana [GHA] (68-93; 68-93) ***
Guinea-Bissau [GNB] (77-81; 77-81)
Kenya [KEN] (84-85; 75-85)
Lesotho [LSO] (81-90; 81-90)
Liberia [LBR] (71-86; 71-86)
Madagascar [MDG] (77-95; 77-95)
Malawi [MWI] (76-97; 68-97)
Mali [MLI] (70-90; 70-90)
Mauritius [MUS] (68-97; 68-97) ***
Mozambique [MOZ] (87-89; 87-89)
Niger [NER] (68-88; 68-88)
Nigeria [NGA] (68-95; 68-95) ***
Reunion [REU] (68-82; 68-82)
Rwanda [RWA] (76-91; 76-91)
Senegal [SEN] (68-92; 68-92) ***
Seychelles [SYC] (68-91; 68-91) ***
Sierra Leone [SLE] (68-96; 68-96) ***
Somalia [SOM] (69-79; 69-79)
South Africa [ZAF] (93-94; 93-94)
Sudan [SDN] (68-97; 68-97) ***
Swaziland [SWZ] (75-93; 75-93)
Tanzania [TZA] (68-72; 68-72)
Togo [TGO] (68-93; 68-93) ***
Uganda [UGA] (93; 93)
Zambia [ZMB] (68-96; 68-96) ***
Zimbabwe [ZWE] (80-87; 80-87)

North Africa & the Middle East (14)

Algeria [DZA] (68-92; 68-92) ***
Bahrain [BHR] (78-98; 78-98)
Cyprus [CYP] (68-97; 68-97) ***
Egypt [EGY] (94-95; 87-95)
Iran [IRN] (69-86; 69-86)
Israel [ISR] (68-81; 68-81)
Jordan [JOR] (68-95; 68-95) ***
Kuwait [KWT] (75; 75)
Morocco [MAR] (68-79; 68-79)
Qatar [QAT] (79-81; 79-81)
Syria [SYR] (68-95; 68-95) ***
Tunisia [TUN] (68-94; 68-97) ***
Turkey [TUR] (75-94; 75-94)
Yemen [YEM] (69-96; 69-96) ***

South & Central America (20)

Argentina [ARG] (68-95; 68-95) ***
Belize [BLZ] (80-95; 80-95)
Bolivia [BOL] (73-97; 73-97) ***
Brazil [BRA] (71-87; 71-87)
Chile [CHL] (69-86; 69-86)
Colombia [COL] (81-89; 81-89)
Costa Rica [CRI] (68-95; 68-95) ***
Ecuador [ECU] (76-79; 76-79)
El Salvador [SLV] (71-97; 68-97) ***
Guatemala [GTM] (68-94; 68-94) ***
Guyana [GUY] (68-97; 68-97) ***
Honduras [HND] (68-97; 68-97) ***
Mexico [MEX] (68-97; 68-97) ***
Nicaragua [NIC] (68-96; 68-96) ***
Panama [PAN] (69-79; 69-79)
Paraguay [PRY] (68-82; 68-82)
Peru [PER] (68-98; 68-98) ***
Suriname [SUR] (70-96; 68-96) ***
Uruguay [URY] (69-95; 69-95) ***
Venezuela [VEN] (68-90; 68-90) ***

North America & the Caribbean (14)

Bahamas, The [BHS] (68-91; 68-90) ***
Barbados [BRB] (68-95; 68-95) ***
Canada [CAN] (68-85; 68-85)
Dominica [DMA] (--; 71)
Dominican Rep. [DOM] (72-97; 72-97) ***
Grenada [GRD] (76-95; 76-95)
Haiti [HTI] (68-88; 68-88)
Jamaica [JAM] (68-79; 68-79)
Puerto Rico [PRI] (68-97; 68-96) ***
St. Kitts & Nevis [KNA] (79-85; 79-85)
St. Lucia [LCA] (68-91; 68-91) ***
St. Vincent [VCT] (74-98; 74-98)
Trinidad & Tobago [TTO] (68-96; 68-96) ***
United States [USA] (68-97; 68-97) ***

Pacific Basin (17)

Australia [AUS] (68-96; 68-96) ***
China [CHN] (90-97; 90-97)
Fiji [FJI] (68-88; 68-88)
Hong Kong, China [HKG] (68-98; 68-98) ***
Indonesia [IDN] (69-92; 69-92) ***
Japan [JPN] (72-96; 68-96) ***
Korea [KOR] (70-97; 76-97)
Malaysia [MYS] (68-95; 68-95) ***
New Zealand [NZL] (68-91; 68-91) ***
Papua New Guinea [PNG] (82-96; 79-96)
Philippines [PHL] (72-95; 72-95) ***
Samoa [WSM] (68-79; 68-79)
Singapore [SGP] (68-96; 68-96) ***
Solomon Islands [SLB] (86; 86)
Taiwan [TWN] (68-70; 68-70)
Thailand [THA] (76-95; 76-95)
Tonga [TON] (75-85; 75-85)

South Asia (6)

Bangladesh [BGD] (69-97; 69-97) ***
India [IND] (68-97; 68-97) ***
Myanmar [MMR] (70-98; 70-98) ***
Nepal [NPL] (75-90; 75-81)
Pakistan [PAK] (68-81; 68-83)
Sri Lanka [LKA] (77-97; 68-95)

Eastern Europe (7)

Bulgaria [BGR] (90; 90)
Czechoslovakia [CSA] (69-91; 69-91);
from 1992-1997, Czech Rep. [CZE] ***
Hungary [HUN] (70-97; 70-97) ***
Romania [ROM] (73-97; 73-97) ***
Russian Fed. [RUS] (88-95; 88-95)
Slovakia [SVK] (95-98; 95-98)
Yugoslavia [YGA] (85-92; 83-92)

Western Europe (19)

Austria [AUT] (68-97; 68-97) ***
Belgium [BEL] (68-98; 68-98) ***
Denmark [DNK] (77-92; 77-92)
Finland [FIN] (68-94; 68-95) ***
France [FRA] (75; 68-75)
Germany, Fed. Rep. [DEU] (68-98; 68-98) ***
Greece [GRC] (68-82; 68-82)
Iceland [ISL] (68-96; 68-96) ***
Ireland [IRL] (68-82; 68-84)
Italy [ITA] (68-97; 68-97) ***
Luxembourg [LYX] (95; 77-95)
Malta [MLT] (68-71; 68-85)
Netherlands [NLD] (68-90; 68-90) ***
Norway [NOR] (75-97; 68-97)
Portugal [PRT] (69-94; 69-94) ***
Spain [ESP] (68-76; 68-76)
Sweden [SWE] (68-95; 68-95) ***
Switzerland [CHE] (68-82; 68-82)
United Kingdom [GBR] (68-98; 68-98) ***

Summary:

Total No. of Countries = 139
Hypothetical Maximum Number of Skilled or Unskilled Wage Data Points = 4309

Actual No. of Unskilled Wage Data Points = 2707 (62.8% coverage)
Actual No. of Skilled Wage Data Points = 2818 (65.4% coverage)

Notes: See Section 3 for details on the derivation of these wage series. From 1992, data from the Czech Republic is linked to data from Czechoslovakia. This wage series should be strictly interpreted as pertaining only to the Czech Republic (excluding Slovakia), assuming that labor markets within Czechoslovakia were well-integrated before its break-up in 1992.

Table 2a
Summary Statistics of Unskilled Occupational Wages (1968-1998)

Statistics that follow refer to the raw un-interpolated real wage data. All wages are deflated by a PPP index for consumption goods and are in 1990 US$.

Occupation Name and Code Number in *October Inquiry*	No. of Observations	5th percentile	Median	95th percentile
Thread and yarn spinner, Textiles industry (#25)	1122	0.61	2.25	7.72
Sewing-machine operator, manufacture of wearing apparel, excluding footwear (#30)	1411	0.61	2.29	7.36
Laborer, printing, publishing and allied industries (#51)	1338	0.44	1.93	8.95
Laborer, manufacture of industrial chemicals and other chemical products (#56/#59) **	1127	0.52	2.13	9.04
Laborer, manufacture of machinery except electrical (#70)	985	0.51	2.55	9.09
Laborer, electric light and power (#80)	1264	0.48	2.28	10.49
Laborer, construction industry (#90)	1574	0.40	2.08	9.58

****** Prior to 1983, only one figure was reported in the *October Inquiry* for the wages of unskilled laborers in the manufacture of chemicals. Occupation #56 (unskilled laborer in the manufacture of industrial chemicals) and occupation #59 (unskilled laborer in the manufacture of other chemical products) were introduced when the scope of the *Inquiry* was expanded in 1983. For consistency with the pre-1983 years, the OWW reports an average of #56 and #59.

Table 2b
Summary Statistics of Skilled Occupational Wages (1983-1998)

Statistics that follow refer to the raw un-interpolated real wage data. All wages are deflated by a PPP index for consumption goods and are in 1990 US$.

Occupation Name and Code Number in *October Inquiry*	No. of Observations	5^{th} percentile	Median	95^{th} percentile
Chemical engineer, manufacture of industrial chemicals (#52)	357	1.18	7.22	22.26
Power distribution and transmission engineer, electric power and light (#76)	455	1.41	8.49	24.34
Bank Accountant (#129)	503	1.40	9.65	22.29
Computer programmer, insurance industry (#133)	401	1.03	6.36	15.29
Government executive official, public administration (#139)	378	1.08	8.64	26.87
Mathematics teacher (third level), education services (#145)	464	1.10	9.51	23.30
General physician, medical and dental services (#152)	539	1.17	10.19	24.57

Table 2c
Summary Statistics of Fairly Skilled Occupational Wages (1968-1998)

Occupation Name and Code Number in *October Inquiry*	No. of Observations	5th percentile	Median	95th percentile
Chemistry technician / Laboratory Assistant, manufacture of industrial chemicals (#53) **	893	1.04	4.27	12.05
Occupational Health Nurse, Iron and Steel Basic Industries (#61)	577	1.02	4.59	13.90
Machinery Fitter-Assembler, manufacture of machinery except electrical (#69)	1053	0.77	3.79	10.79
Building electrician, construction (#81)	1523	0.77	3.41	11.37
Stenographer-typist, wholesale trade / grocery (#91)	1242	1.02	3.81	9.21
Stored records clerk, wholesale trade / grocery (#92)	1280	0.84	3.29	8.65
Salesperson, retail trade / grocery (#96)	1399	0.72	2.49	7.64
Bank teller, banks (#131)	1374	1.02	4.56	10.95
Book-keeping machine operator (#132)	1205	1.00	4.33	10.80
Automobile mechanic, repair of motor vehicles (#159)	1482	0.96	3.30	10.38

** Wages reported prior to 1983 for laboratory assistants were matched with those reported for chemistry technicians post-1983, given the good match in job descriptions between the two occupations.

Table 3a
Summary Results from Fixed Effects Regressions of Unskilled Occupational Wages (1968-1998)

Dependent variable: occupation listed in the leftmost column

Occn. No.	#25	#30	#51	#56/#59	#70	#80	#90
#25	NA	0.87 (55.85) *** R^2 = 0.89 SEP: 0.085; 0.194	0.82 (44.06) *** R^2 = 0.86 SEP: 0.122; 0.208	0.74 (39.50) *** R^2 = 0.87 SEP: 0.125; 0.189	0.95 (53.26) *** R^2 = 0.86 SEP: 0.126; 0.159	0.52 (27.65) *** R^2 = 0.81 SEP: 0.122; 0.253	0.71 (35.91) *** R^2 = 0.83 SEP: 0.109; 0.249
#30	0.77 (55.85) *** R^2 = 0.89 SEP: 0.071; 0.128	NA	0.77 (46.06) *** R^2 = 0.86 SEP: 0.072; 0.153	0.64 (35.95) *** R^2 = 0.84 SEP: 0.083; 0.163	0.87 (53.05) *** R^2 = 0.85 SEP: 0.094; 0.139	0.43 (25.26) *** R^2 = 0.76 SEP: 0.090; 0.196	0.55 (34.26) *** R^2 = 0.82 SEP: 0.088; 0.196
#51	0.70 (44.06) *** R^2 = 0.86 SEP: 0.113; 0.138	0.69 (46.06) *** R^2 = 0.86 SEP: 0.087; 0.144	NA	0.71 (49.76) *** R^2 = 0.91 SEP: 0.096; 0.117	0.82 (45.97) *** R^2 = 0.91 SEP: 0.082; 0.135	0.55 (41.03) *** R^2 = 0.86 SEP: 0.096; 0.137	0.48 (47.21) *** R^2 = 0.85 SEP: 0.095; 0.153
#56/#59	0.73 (39.50) *** R^2 = 0.87 SEP: 0.139; 0.173	0.70 (35.95) *** R^2 = 0.84 SEP: 0.111; 0.214	0.85 (49.76) *** R^2 = 0.91 SEP: 0.094; 0.168	NA	0.91 (45.84) *** R^2 = 0.94 SEP: 0.087; 0.160	0.48 (40.84) *** R^2 = 0.88 SEP: 0.083; 0.203	0.59 (35.46) *** R^2 = 0.88 SEP: 0.085; 0.219
#70	0.74 (53.26) *** R^2 = 0.86 SEP: 0.135; 0.148	0.78 (53.05) *** R^2 = 0.85 SEP: 0.111; 0.180	0.74 (45.97) *** R^2 = 0.91 SEP: 0.082; 0.188	0.69 (45.84) *** R^2 = 0.94 SEP: 0.086; 0.167	NA	0.51 (31.87) *** R^2 = 0.88 SEP: 0.087; 0.241	0.57 (37.46) *** R^2 = 0.88 SEP: 0.099; 0.236
#80	0.71 (27.65) *** R^2 = 0.81 SEP: 0.147; 0.224	0.65 (25.26) *** R^2 = 0.76 SEP: 0.135; 0.246	0.88 (41.03) *** R^2 = 0.86 SEP: 0.113; 0.177	0.80 (30.84) *** R^2 = 0.88 SEP: 0.099; 0.206	0.89 (31.87) *** R^2 = 0.88 SEP: 0.105; 0.222	NA	0.48 (23.70) *** R^2 = 0.79 SEP: 0.117; 0.252
#90	0.64 (35.91) *** R^2 = 0.83 SEP: 0.131; 0.148	0.69 (34.62) *** R^2 = 0.82 SEP: 0.131; 0.185	1.08 (47.21) *** R^2 = 0.85 SEP: 0.116; 0.194	0.53 (35.46) *** R^2 = 0.88 SEP: 0.095; 0.164	0.88 (37.46) *** R^2 = 0.88 SEP: 0.114; 0.177	0.48 (23.70) *** R^2 = 0.79 SEP: 0.106; 0.207	NA

Notes: The first line in each cell is the slope coefficient estimate, followed by absolute t-statistics in parentheses; \ *** indicates slope coefficient is significantly different from 0 at the 1% level. The second row reports the overall R^2 of the fixed effects regressions. The third row in each cell indicates the average standard error of the predictors (SEP), the first figure being for the predictor of out-of-sample country fixed effects, and the second figure referring to the fixed effects linear wage predictor. For more details, please see Section 3.3.

Table 3b
Summary Results from Fixed Effects Regressions of Skilled Occupational Wages (1983-1998)

Dependent variable: occupation listed in the leftmost column

Occcn. No.	#52	#76	#129	#133	#139	#145	#152
#52	NA	0.17 (2.48) ** R^2 = 0.62 SEP: 0.740; 1.543	0.120 (1.73) * R^2 = 0.33 SEP: 0.933; 1.484	0.79 (6.67) *** R^2 = 0.57 SEP: 0.794; 1.284	0.28 (6.17) *** R^2 = 0.52 SEP: 0.968; 1.030	0.24 (3.30) *** R^2 = 0.52 SEP: 0.871; 1.591	0.04 (0.52) R^2 = 0.56 SEP: 0.843; 1.519
#76	0.13 (2.48) ** R^2 = 0.62 SEP: 0.669; 1.121	NA	0.25 (4.93) *** R^2 = 0.60 SEP: 0.643; 1.119	0.40 (4.61) *** R^2 = 0.61 SEP: 0.824; 1.070	0.53 (10.22) *** R^2 = 0.46 SEP: 0.965; 0.898	0.35 (10.31) *** R^2 = 0.67 SEP: 0.632; 0.893	0.44 (9.67) *** R^2 = 0.70 SEP: 0.688; 1.027
#129	0.08 (1.73) * R^2 = 0.33 SEP: 0.789; 0.925	0.22 (4.92) *** R^2 = 0.60 SEP: 0.569; 0.980	NA	0.39 (5.57) *** R^2 = 0.54 SEP: 0.632; 0.801	0.11 (2.01) ** R^2 = 0.34 SEP: 0.868; 1.220	0.26 (5.56) *** R^2 = 0.43 SEP: 0.701; 1.086	0.32 (6.16) *** R^2 = 0.62 SEP: 0.558; 0.989
#133	0.17 (6.67) *** R^2 = 0.57 SEP: 0.500; 0.532	0.14 (4.61) *** R^2 = 0.61 SEP: 0.539; 0.695	0.19 (5.57) *** R^2 = 0.54 SEP: 0.504; 0.650	NA	0.14 (5.46) *** R^2 = 0.64 SEP: 0.502; 0.658	0.24 (10.19) *** R^2 = 0.60 SEP: 0.640; 0.547	0.18 (6.93) *** R^2 = 0.66 SEP: 0.517; 0.644
#139	0.60 (6.17) *** R^2 = 0.52 SEP: 1.256; 1.466	0.49 (10.22) *** R^2 = 0.46 SEP: 1.129; 1.057	0.13 (2.01) ** R^2 = 0.34 SEP: 1.278; 1.587	0.80 (5.46) *** R^2 = 0.64 SEP: 1.229; 1.673	NA	0.46 (7.69) *** R^2 = 0.41 SEP: 1.038; 1.252	0.47 (8.12) *** R^2 = 0.59 SEP: 0.922; 1.170
#145	0.16 (3.30) *** R^2 = 0.52 SEP: 0.839; 1.040	0.58 (10.31) *** R^2 = 0.67 SEP: 0.545; 1.118	0.28 (5.56) *** R^2 = 0.43 SEP: 0.741; 1.125	0.95 (10.19) *** R^2 = 0.60 SEP: 0.824; 1.011	0.32 (7.69) *** R^2 = 0.41 SEP: 0.859; 0.849	NA	0.69 (11.24) *** R^2 = 0.58 SEP: 0.666; 1.132
#152	0.02 (0.52) R^2 = 0.56 SEP: 0.856; 0.826	0.40 (9.67) *** R^2 = 0.70 SEP: 0.621; 0.864	0.24 (6.16) *** R^2 = 0.62 SEP: 0.670; 0.810	0.61 (6.93) *** R^2 = 0.66 SEP: 0.860; 0.984	0.35 (8.12) *** R^2 = 0.59 SEP: 0.838; 0.804	0.31 (11.24) *** R^2 = 0.58 SEP: 0.665; 0.705	NA

Notes:
See notes for table 3a. *, **, and *** respectively indicate that the slope coefficient is significantly different from 0 at the 10%, 5%, and 1% levels.

Table 3c
Regression Results from Proxying for the Skilled Wage Series

Dependent variable: Skilled wage series 1983-1998

Instrument Occn. No.	Slope Coefficient	Overall R^2	Root Mean Square Error	Standard Error of linear predictor (fixed effects regression)	Standard Error of predictor for out-of-sample country fixed effects
#53	0.20 (8.35) ***	0.78	0.306	0.450	0.354
#61	0.24 (13.62) ***	0.75	0.373	0.598	0.414
#69	0.49 (10.70) ***	0.73	0.671	0.561	0.513
#81	0.35 (10.56) ***	0.59	0.647	0.535	0.445
#91	0.44 (8.64) ***	0.78	0.619	0.506	0.473
#92	0.43 (10.02) ***	0.68	0.587	0.552	0.439
#96	0.46 (12.83) ***	0.64	0.598	0.480	0.423
#131	0.36 (10.78) ***	0.79	0.585	0.401	0.424
#132	0.23 (6.78) ***	0.77	0.733	0.381	0.494
#159	0.28 (7.85) ***	0.75	0.863	0.453	0.505

Notes: Absolute t-statistics are in parentheses. *** indicates slope coefficient is significantly different from 0 at the 1% level. For details, please see Section 3.4.

Table 4
Regression Results on Unconditional Beta Convergence (67 countries)

Dependent variable: Growth rate of designated variable (in % per annum)

	(1) Real GDP per worker	(2) Real Skilled Wages	(3) Real Unskilled Wages	(4a) Relative Skilled-Unskilled Wage	(4b) Relative Skilled-Unskilled Wage
Initial Level	0.0000249 (0.723)	0.0557 * (1.960)	-0.149 (1.527)	-0.197 *** (2.831)	-0.350 *** (3.637)
Constant	0.782 ** (2.181)	-0.846 *** (2.835)	0.455 (1.219)	0.488 (1.358)	1.044 ** (2.433)
Adjusted R^2	-0.0075	0.0413	0.0198	0.0961	0.160
Root Mean Square Error	1.840	1.193	2.032	1.604	1.559
Convergence "Half-Life"	Divergence	Divergence	63 years Gap: 7.41 US$	42 years Gap: 8.47 US$	33 years Gap: 5.92 US$

Notes: Absolute t-statistics are in parentheses; *, **, and *** denote significance at the 10%, 5% and 1% level respectively. The number of observations is 67, except for (4b) where two outliers (BDI and NGA) are removed. The convergence "half-life" is the number of years required to halve the gap (reported in the ``Gap'' row) between the 5th and 95th percentile of the 1994 distribution of the relevant income or wage variable. See Section 4.2 for details.

Table 5
Regression Results on Unconditional Beta Convergence (34 Open Economies)

Dependent variable: Growth Rate of Designated Variable (in % per annum)

	(1) Real GDP per worker	(2) Real Skilled Wages	(3) Real Unskilled Wages	(4) Relative Skilled-Unskilled Wage
Initial Level	-0.0000451 (1.069)	0.0319 (1.128)	-0.300 *** (2.961)	-0.448 *** (4.131)
Constant	2.140 *** (3.894)	-0.351 (1.024)	1.843 *** (3.840)	1.072 ** (2.154)
Adjusted R^2	0.0043	0.0082	0.191	0.328
Root Mean Square Error	1.775	0.829	1.777	1.363
Convergence "Half-Life"	53 years Gap: 29245 US$	Divergence	29 years Gap: 7.91 US$	33 years Gap: 4.65 US$

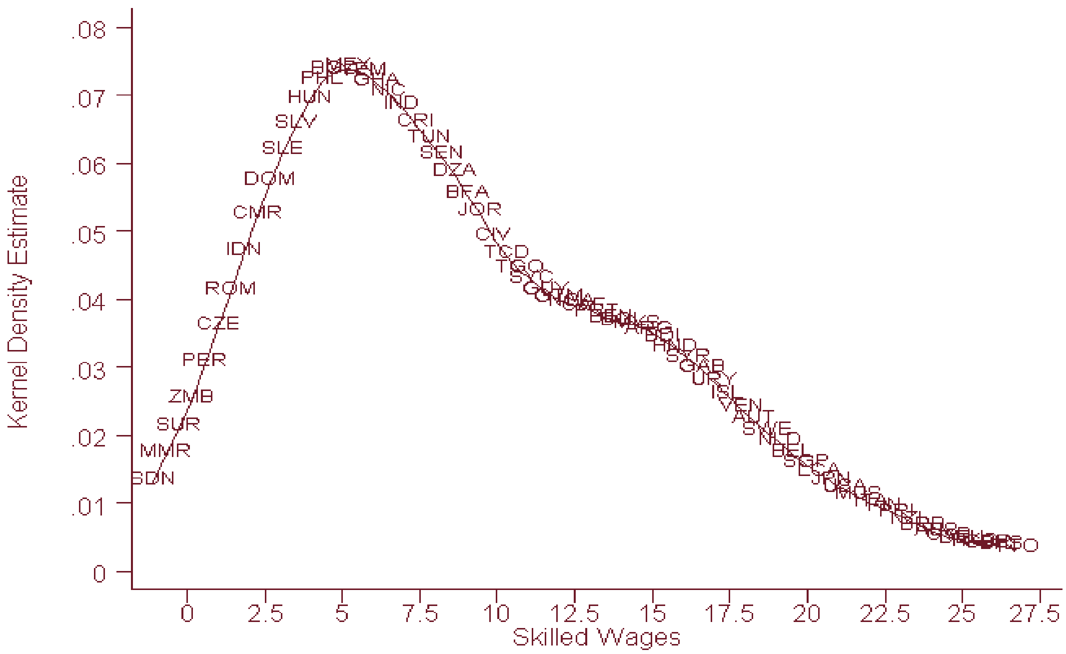
Notes: See Table 4. Countries are classified as open or closed according to the Sachs and Warner (1995) criterion.

Table 6
Comparison of Wage and Aggregate Income Growth Rates for Open versus Closed Economies
(% per annum)

Annual Growth Rate of:		Open	Open (OECD only)	Open (non-OECD only)	Closed
Skilled Real Wages	70-82:	-0.15	+0.29	-0.43	-0.87
	83-94:	+0.08	+0.09	+0.08	-0.47
	70-94:	+0.00	+0.19	-0.12	-0.71
Unskilled Real Wages	70-82:	+0.72	+1.14	+0.46	-0.41
	83-94:	+0.64	+0.64	+0.64	-1.39
	70-94:	+0.74	+0.92	+0.64	-0.86
Skilled/Unskilled Wage	70-82:	-0.87	-0.85	-0.88	-0.46
	83-94:	-0.56	-0.55	-0.57	+0.93
	70-94:	-0.74	-0.73	-0.76	+0.15
Real GDP per Worker	70-82:	+1.95	+1.71	+2.09	+1.45
	83-94:	+1.54	+1.82	+1.36	-0.86
	70-94:	+1.65	+1.74	+1.59	+0.20

Figure 1
The Worldwide Distribution of Hourly Real Wages, 1994

a: Skilled Wages (1990 US$)



b: Unskilled Wages (1990 US$)

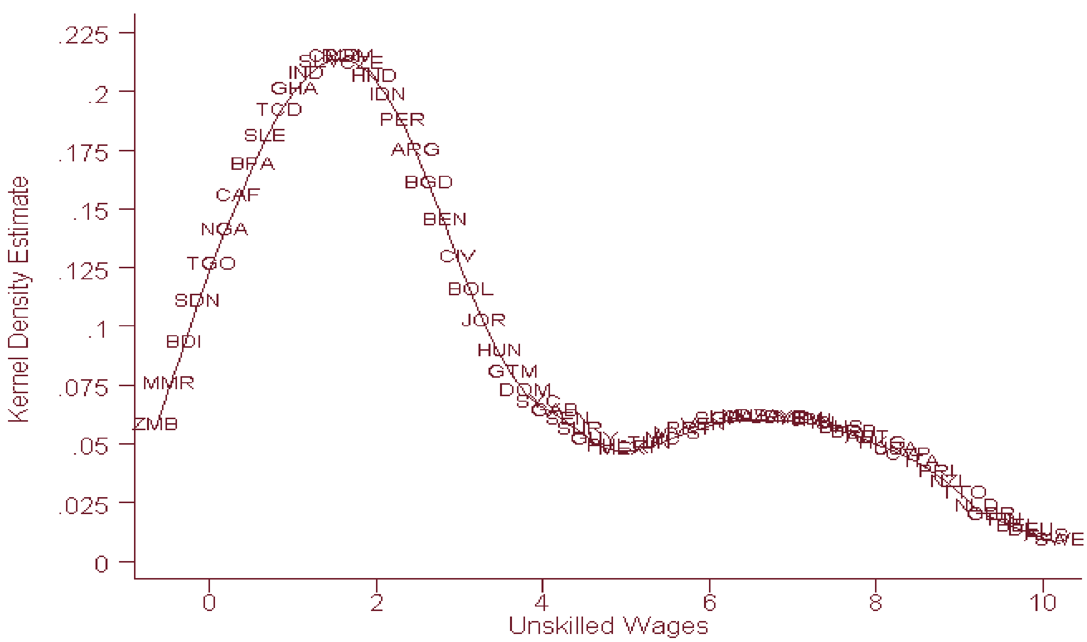
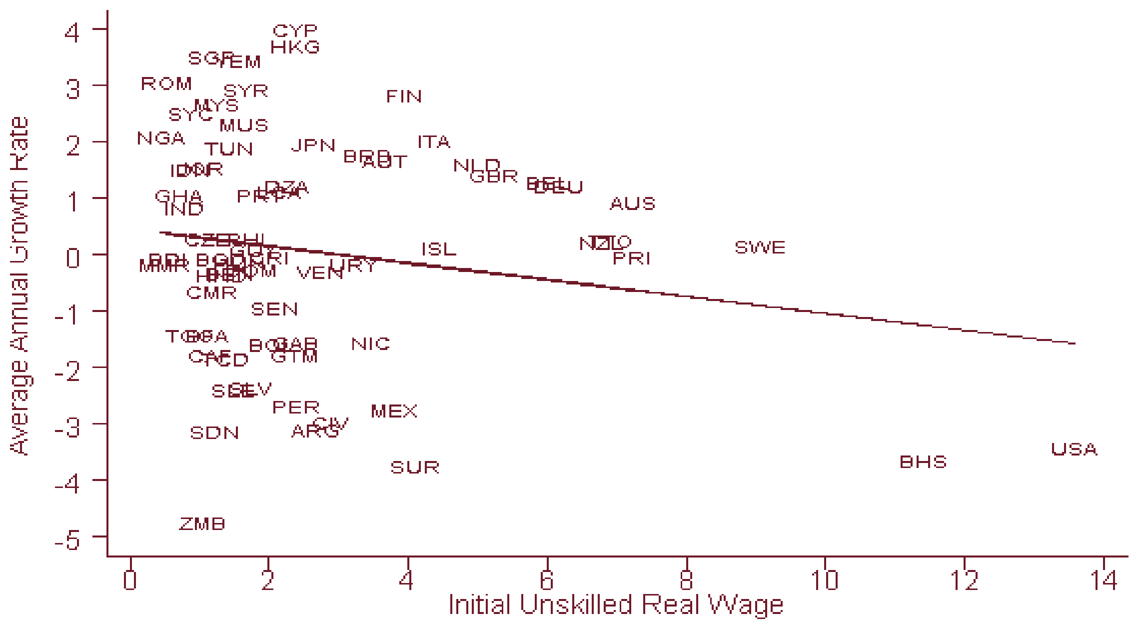


Figure 2
Unconditional Beta Convergence (Sample of 67 countries)

a: Unskilled Real Wage, 1990 US$



b: Skilled-Unskilled Wage Ratio

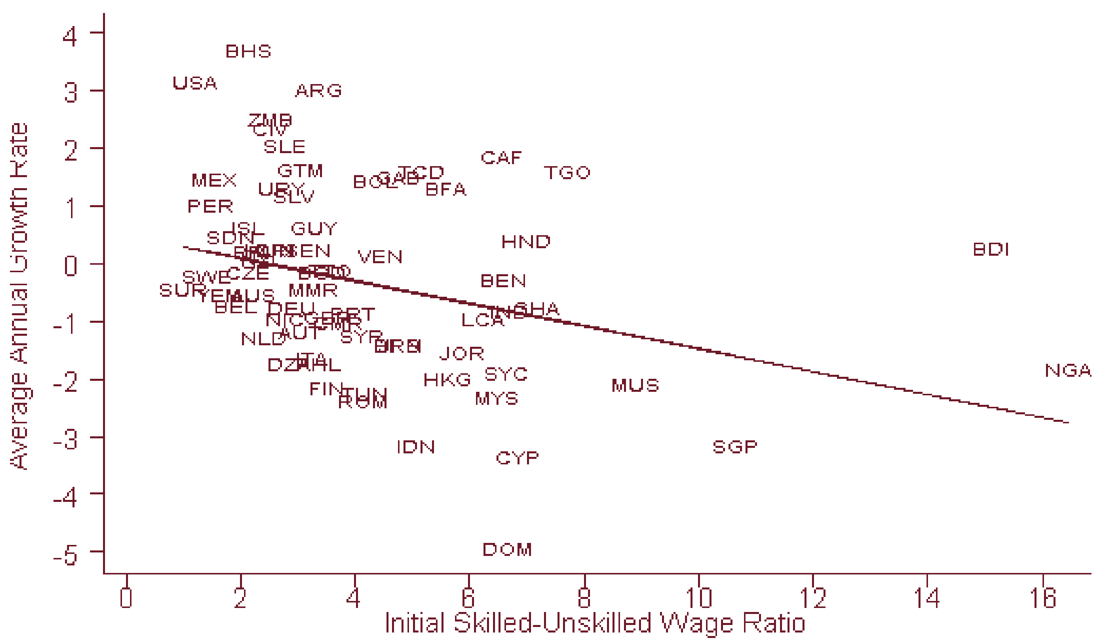
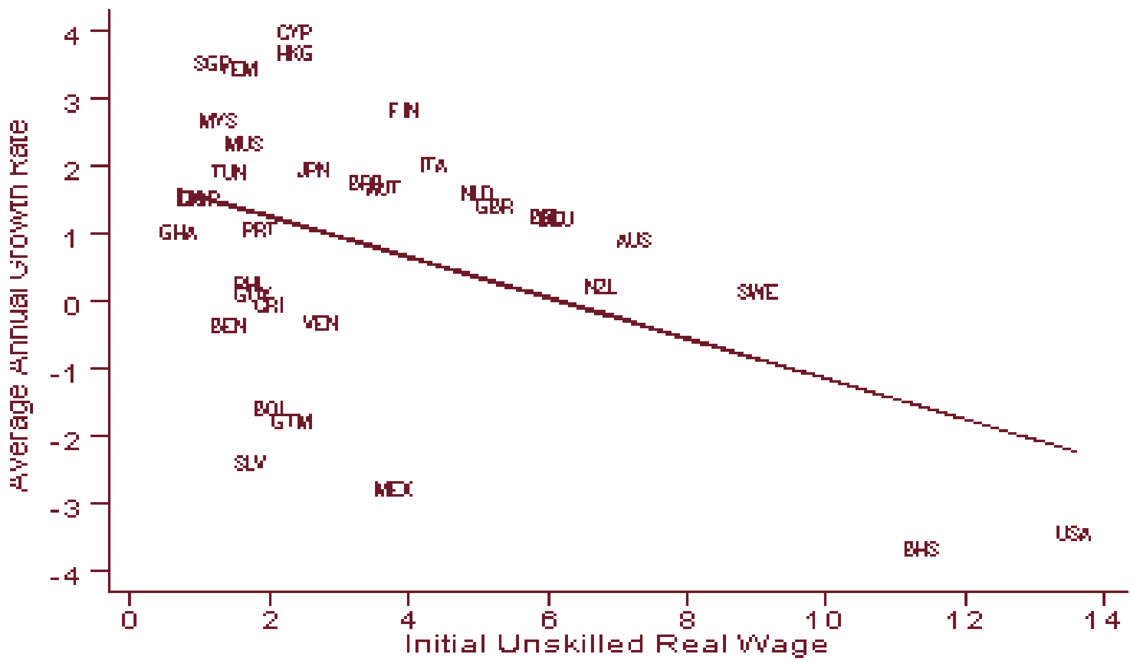


Figure 3
Unconditional Beta Convergence (Convergence Club of 34 Open Economies)

a: Unskilled Real Wage, 1990 US$



b: Skilled-Unskilled Wage Ratio