GENDER WAGE GAPS IN CHILE. A LONG TERM VIEW: 1958-1990

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ABSTRACT

Using a statistical technique to correct for the presence of the selectivity bias, this paper estimates gender wage gaps in Chile for the period 1958-1990. Gender wage gaps are decomposed into an "endowment" effect --associated with different stocks of human capital-- and a "discrimination" effect --associated with different returns. The paper argues that a tight labor market would increase the cost of discriminatory practices, fostering the reduction in wage gaps. The evidence reports that observed changes in gender wage gaps largely respond to the economic cycle, suggesting that anti-discriminatory policies must be considered only as a complement to policies aimed at attaining sustainable economic growth.

SÍNTESIS

Usando una técnica estadística para corregir la presencia del sesgo de selección, este trabajo estima las diferencias de salario por género en Chile para el período 1958-1990. Las diferencias en salarios por género se descomponen en un efecto de "dotación" --asociado a diferentes niveles de capital humano-- y en un efecto de discriminación --asociado a diferentes retornos. El trabajo plantea que los mercados laborales con bajo desempleo aumentarían el costo de las prácticas discriminatorias, lo que estimula la disminución de las diferencias en salarios. La evidencia demuestra que los cambios observados en las diferencias en salarios por género responden en gran medida al ciclo económico, sugiriendo que las políticas anti-discriminatorias deben ser considerados sólo como un completo de las políticas orientadas a lograr un desarrollo económico sustentable.

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1. INTRODUCTION

The need to adopt anti-discrimination measures aimed at equalizing employment opportunities for different labor force groups is a relatively recent policy concern in developing countries (LDCs). Long-lived and important labor market discrimination against certain ethnic groups and women are explained by cultural and historical factors which inspire regulations. Discrimination against women is particularly worrisome as it may constitute a key factor limiting the participation of this group in the labor force and the attainment of more social and economic progress. Even though the reduction of the degree of gender discrimination is a politically accepted principle, customary policies have proved to be insufficient to introduce a more egalitarian access to the labor market. Although the practice of general anti-discriminatory policies has been on the rise in industrial countries (Leonard, 1984), in LDCs the issue has been largely restricted to both the enactment of additional protective polices and the introduction of partial deregulation allowing the access of women to some types of jobs.

In attempting to reduce the importance of existing gender discrimination in LDCs, the attainment of higher and sustained economic growth can be considered just as important as adopting deregulatory policies addressing the labor market. Low growth rates and the existence of a substantial labor surplus make the practice of wage and employment discrimination against definite groups far easier. Within this context, the introduction of deregulatory policies to reduce the extent of the gender bias may only partly achieve more competitive labor markets outcomes, if such policies are not accompanied by economic growth and a less

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tight labor market. The switch in the policy discussion would, thus, focus more on growth promotion and general measures aimed at deregulating the labor market, instead of focusing on specific anti-discriminatory regulations.

Greater emphasis has been devoted in the literature to microeconomic empirical analyses of gender wage gaps, based upon cross-sectional studies. Time trend analyses are, however, important to assess the likely effect of macroeconomic outcomes and the relative importance of growth vis à vis deregulation in confronting the undesirable effects of discrimination. In doing so, to accurately measure gender discrimination, beyond simple gender wage gaps, becomes a critical need.

This paper estimates a time series of gender wage gaps for the period 1950-1990 in Chile. The analysis focuses on the existence of two different explanatory factors of observed gender wage gaps: pure discrimination generated in the labor market, and discrimination originated in the access to education and human capital formation. Using a statistical technique to correct for the selectivity bias, the paper estimates gender wage gaps and decomposes them into an "endowment" effect --associated with different stocks of human capital-- and a "discrimination" effect. We argue that observed changes in gender wage gaps respond to the economic cycle, whereas anti-discriminatory policies are only a complement to policies aimed at attaining sustainable growth. In other words, a tight labor market would be associated with an increase in the cost of discriminatory practices that foster decreases in wage gaps thereby implying that greater emphasis must be lent to growth and macroeconomic factors.

The paper contains five sections. The second section discusses the methodology. The third section analyzes the estimated gender wage gaps and discusses the importance of the selectivity bias correction in the Chilean case. The fourth section uses a model in which GDP and other labor market variables account for the evolution of the estimated gender wage gaps associated with discrimination over time. In section 5 we conclude.

2. GENDER WAGE GAPS AS A POLICY ISSUE

The persistence of existing wage gaps across different labor force groups, for labor with similar attributes, have become a growing concern in most industrial countries. This concern has been related mostly to existing labor market practices and distortions. Allegations concerning systematic discriminatory practices have led to considering the need for suitable regulations in relation to employment and wage practices. Research in this area has provided background material for a growing anti-discriminatory legislation, especially in the USA where discrimination against women has been usually analyzed along the same lines as that of racial discrimination. However, the success of this anti-discriminatory legislation in attaining lower wage gaps has been largely controversial. Concurrently, alternative empirical models have not been convincing enough in explaining the persistence of wage gaps throughout time despite a more effective equalization of human capital endowments.[1]

In the case of Latin America, the literature that has analyzed the presence of gender discrimination has been all but abundant. An exception is the collection of studies edited by Psacharopoulos and Tzannatos (1991) covering Latin American countries and which concludes that the extent of gender discrimination is important in the countries studied, and that the degree of the selectivity bias is high due to the presence of relatively low female labor force participation (LFP) rates. These studies, however, have neither addressed the issues surrounding the behavior of the estimated "potential" gender discrimination throughout time nor considered the effect of the economic cycle on gender wage gaps. To the contrary, this literature has heavily relied upon cross sectional studies, focusing primarily on microeconomic aspects of the wage determination process. In the case of Chile, the study by Gill (1991) has arrived at similar conclusions as those of Psacharopoulos and Tzannatos, even though his empirical analysis did not control for changes in hours worked, which is a crucial variable in explaining the behavior of total labor earnings. In another study on Chile, Paredes (1982) concluded that measured gender wage gaps reached an equivalent to 50 percent of males' wages in 1969, but that they declined towards 1978. Although a great proportion of measured gaps is explained by differences in human capital endowments between males and females, observed differences in estimated parameters, which are likely to be associated with discrimination, were also found to be very important.

$$E\ (W_M/\ I > \mu_p/\sigma_p)\ =\ \alpha X\ -\ \frac{\sigma_{1P}}{\sigma_P}\lambda$$

Though crucial in the case of the population that is presumably discriminated, most empirical analyses on existing wage gaps do not correct for the presence of any selectivity bias. Given the existence of lower LFP rates for females, the presence of the selectivity bias may introduce an overestimation in relation to the effect of independent variables —particularly those associated with human capital— on women's wages. This overestimation may be higher than that associated with the presence of the selectivity bias in the case of males, due precisely to the differences in LFPs for both groups. As stated, not only the female LFP is normally much lower than in the case of males, but also open

[1] Becker (1971), Mincer and Polacheck (1978) and Corcoram and Duncan (1979) the classics in analyzing theoretical issues. Lazear (1989) presents a more policy-oriented view of this issue.

unemployment is notably higher for the former group. The empirical overestimation of predicted wages would, in turn, derive from the acceptable assumption that the population included in the sample has relatively more human capital than the population actually excluded. This would imply that average earnings for the average person in the sample (the predicted population wage) would be higher. Therefore, measures of gender wage gaps which do not account for the selectivity bias problem may lead to seriously underestimating the effect of gender discrimination on wages in any given period. Futhermore, as the female LFP has increased over time, the failure to correct for the selectivity bias problem would conceal a recovery of the relative wage for women, as has been the case in the USA over the past two decades (Smith and Ward, 1989).

2.1. Methodology

To empirically estimate corrected gender wage gaps, the following steps were taken. First, the sample was broken up into the male and the female labor force. Second, a standard (Mincer type) earnings function was estimated for those employed (with observed incomes) in both groups. Third, the OLS estimates were corrected for the selectivity bias, following Heckman's (1979) methodology. This correction implies the inclusion in the OLS equation of the predicted value of the inverse of the Mill's ratio, derived from a "participation" equation which accounts for the probability to be observed with a positive wage. Fourth, the fitted wage from estimated corrected equations for both males and females was obtained. Fifth, the fitted wage for females was calculated, using females' observed average human capital and estimated coefficients. Finally, the potential discrimination against women was calculated by means of the existing gap between the actual fitted wage for females and the value that would result from estimating females' human capital based on observed males' coefficients.

The data comes from the quarterly labor force survey carried out by the Department of Economics of the University of Chile for the Greater Santiago Area. The information of this study corresponds to the months of June over the period 1958-1990. The total sample in each individual year is of about 7,000 individuals. Workers in the category of "domestic services" have been excluded for the purposes of this study, since in their case the information on hours worked and wages is deemed to be unreliable.[2] This data source has been used in many other studies aimed at measuring the relationship between earnings and human capital (see, for instance, Corbo and Stelcner, 1981, Uthoff, 1981, Riveros, 1990, and Basch and Paredes, 1992) and has yielded very consistent results. The sample has been periodically updated to reflect the growth of the Santiago Area,

[2] It is important to report that this group includes a disproportionately large group of female workers.

and the questionnaire has been modified only in minor aspects, making it possible to rely on fairly comparable information over time.

When empirically measuring gender wage gaps, a distinction has been made between the "pure" discrimination and the observed "market" wage gap (associated with differences in human capital endowments). "Pure" discrimination is associated with the existence of different rates of return to human capital by gender, which corresponds to a different method of remunerating similar human capital levels.[3] A Mincer-type earnings function is specified, in which hourly wages are explained by years of formal schooling and potential market experience is defined as age minus schooling minus six.[4] This specification is estimated separately for males and females. To correct for the selectivity bias problem, the Heckman (1979) approach was used. Accordingly, a "participation equation"[5] was estimated as a first step, using a Probit procedure. The participation equation allows for estimating the probability of being included in the labor force as a function of a set of explanatory variables. The predicted probability of being included in the labor force allows us to create the variable ***lambda*** (the inverse of the Mill's ratio), which is included as an additional independent variable in the OLS wage regression.

Figure 1 illustrates the problem arising from the presence of the selectivity bias, on the assumption that it only affects the female population. In this Figure we also assume that experience is the only human capital variable affecting the log of hourly wages, and that the random error term is distributed normally with constant variance. The OLS estimate βx corresponds to the whole population, implying that β is an unbiased estimate of the population parameter ß. The selectivity bias problem results from considering a non-representative sample of the population, which excludes individuals that are not actually working and earning an income. It is, for instance, highly likely that a significant proportion of women is excluded from the sample, and that the excluded population is not randomly distributed across human capital levels. For instance, the proportion of population excluded from the sample decreases at higher human capital levels. Hence, the OLS line --when the excluded population is not a random sample of the population-- will be b*X instead. In the case depicted in Figure 1, the OLS line, or the observed wage for the working population overestimates the predicted

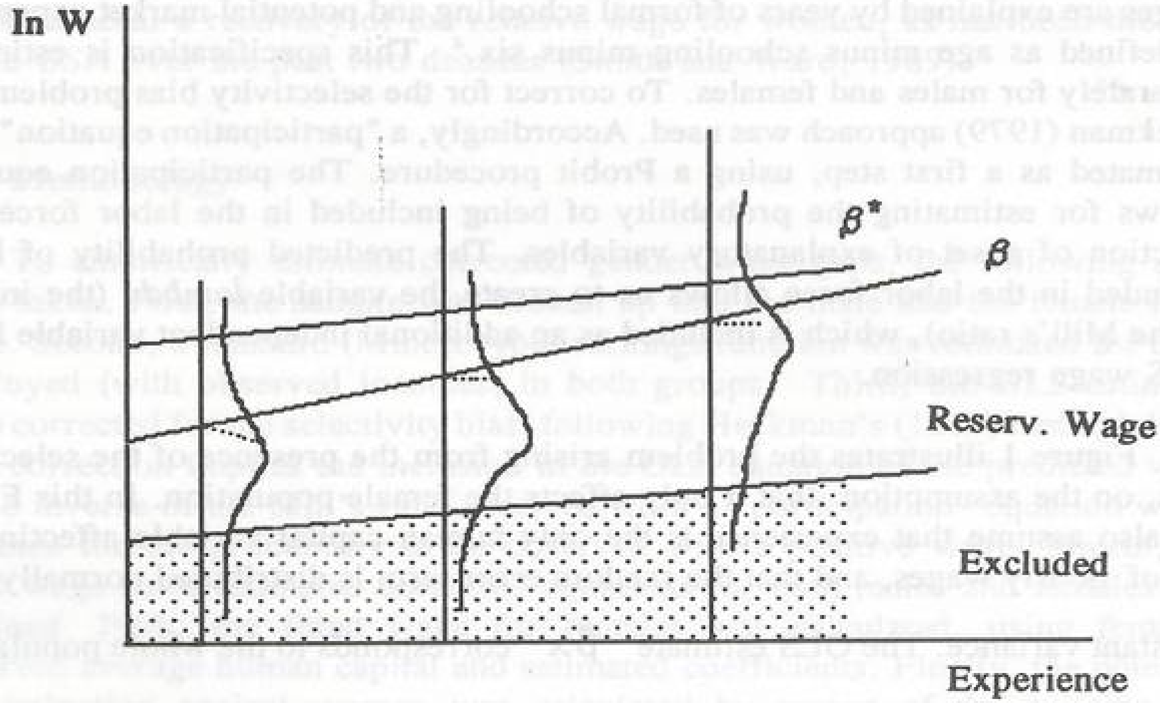
[3] See, to this respect, Psacharoupoulos and Tzannatos (1991).

[4] To account for the likely different effects of the potential experience variable, an estimation of the marital status and the number of children was included; surprisingly, the results obtained when including those variables do not differ from the estimations which do not consider them, so no further analysis is carried out in this connection. For a detailed analysis on this point, see Mincer and Polackeck (1978), and Malkiel and Malkiel (1973).

[5] Rather than labor force participation, this was a variable which measured the probability to be observed with a positive wage.

population wage in different degrees for the different experience levels, since it has been assumed that the people excluded have abnormally low non-observable characteristics or negative errors.

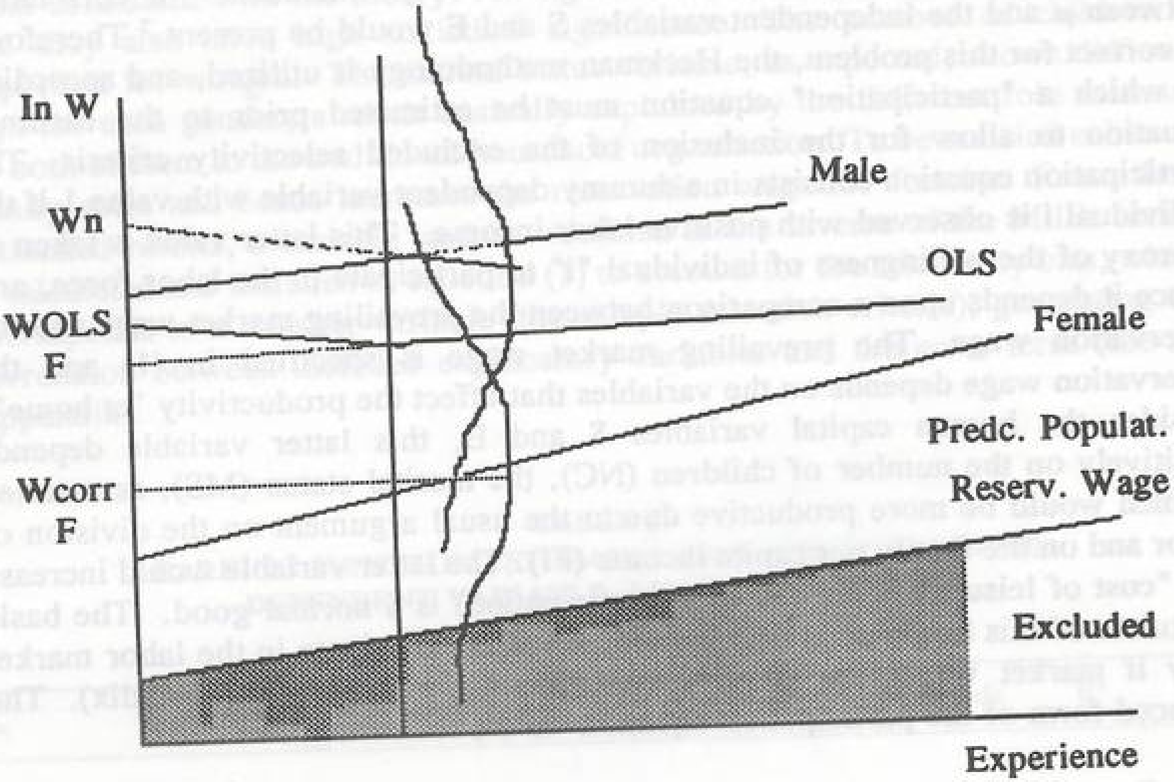
FIGURE 1



In figure 2 the male sample is also included to illustrate the case in which the (OLS) difference underestimates the true male-female wage gap for the whole population. Assuming that the male wage structure is not "discriminatory"[6] for males with an experience equal to Xo, the predicted population wage is also the OLS fit W^{PP}_{M}. For females the predicted population wage is W^{PP}_{F}, but the OLS model would predict a wage W_F OLS and, consequently, a lower "empirical" difference ($W_{PP}^{M}-W_{OLS}^{F}$) than the "true" one ($W_{PP}^{M}-W_{PP}^{F}$)

[6] See, to this respect, Cotton (1988), who suggests that the non-discriminatory structure is somewhere in between that of the male and female clusters.

FIGURE 2



It also appears from figure 2 that observed wage gaps increase with experience. This result derives from the fact that the female participation rate increases with experience and hence a negative correlation between experience and μ is created. In turn, this would generate a negative bias on the return to experience.

3. ESTIMATION AND RESULTS

The empirical analysis aims at estimating the gender wage gap associated with differences in rates of return to human capital. This gap is interpreted as an empirical measurement of wage discrimination. The empirical specification of the wage equation is:

$$\ln w = \alpha_0 + \alpha_1 S + \alpha_2 E + \alpha_3 E_2 + \alpha_4 (S*E) + \mu \quad (1)$$

where w is hourly wages, S and E denote years of schooling and potential experience, respectively, the α's are parameters and μ is a random term.

As mentioned in section 2, when the whole population is considered, μ is well behaved in the sense that it is normally distributed with zero mean, a constant variance and not correlated with S or E. In a sample characterized by the selectivity bias --in which the population "excluded" from the sample is not properly represented by the population effectively "included"-- a correlation between μ and the independent variables S and E would be present.[7] Therefore, to correct for this problem, the Heckman methodology is utilized, and according to which a "participation" equation must be estimated prior to the earnings equation to allow for the inclusion of the excluded selectivity criteria. The participation equation consists in a dummy dependent variable with value 1 if the individual i is observed with positive labor income. This latter value is taken as a proxy of the willingness of individual "i" to participate in the labor force, and hence it depends upon a comparison between the prevailing market wage and the reservation wage. The prevailing market wage is specified in (1) and the reservation wage depends on the variables that affect the productivity "at home". Besides the human capital variables S and E, this latter variable depends positively on the number of children (NC), the marital status (MS), as married women would be more productive due to the usual argument on the division of labor and on the family per capita income (FI). The latter variable would increase the "cost of leisure", if leisure (non-participation) is a normal good. The basic postulate of this behavior is that individual i will participate in the labor market only if market wages are above reservation wages (see the Appendix). The reduced form of the participation equation is the following:

$$\text{Part} = \beta_0 + \beta_1 S + \beta_2 E + \beta_3 E^2 + \beta_4 (S * E) + \beta_5 NC + \beta_6 MS + \beta_7 FI \qquad (2)$$

Table 1 shows the results for the participation equation obtained for the year 1990. This is done to illustrate the empirical results and the estimating technique, as similar results were obtained for each of the cross sections in the period 1958-1990. As stated, the focus of this analysis is not on cross-sectional results, but rather on the time series of estimated gender wage gaps.

TABLE 1
LABOR FORCE PARTICIPATION EQUATION (PROBIT)
DEPENDENT VARIABLE: PARTICIPATE = 1
(1990)

	β_0	β_1	β_2	β_3	β_5	β_5	β_6	β_7	R^2
Men	-.16 (-4.23)	.43 (13.94)	.060 (29.43)	-.0009 (-36.3)	-.0014 (-13.88)	-.006 (-.84)	.17 (8.27)	-.00015 (-9.69)	.50
Women	-.65 (-15.9)	.076 (23.3)	.038 (19.4)	-.0004 (-18.5)	-.0015 (15.07)	-.015 (-2.50)	-.14 (-8.66)	-.00007 (-5.61)	.21

Each of the parameters corresponds to the specification (2). t ratios are between brackets.

[7] As explained above, when human capital increases, more and more women with negative market errors are included, since it is possible for them to get a market wage higher than the reservation wage.

In general, results of the participation equations for male and female are quite consistent with the theory. All signs are the expected ones for each group and they also have high statistical significance. The number of children is important for women. The marital status variable, as expected, shows different sign for each gender, a result basically explained by the different role assumed by both of them in the traditional household organization. To be married enhances family roles and hence increases the reservation wage for females. From these estimated results, a variable lambda --defined as the inverse of the Mill's ratio-- is calculated and included in equation (1) to account for the selectivity bias, which corresponds to a missing variable (the sample selection criteria), generating the correlation between included explanatory variables and the error term (see the Appendix).

TABLE 2
EARNING FUNCTION ESTIMATES (OLS AND CORRECTED)
DEPENDENT VARIABLE, LN HOURLY WAGES
(1990)

α_0	α_1	α_2	α_3	α_4	Lambda	R^2
Men						
OLS	2.10	0.24	0.10	-.001	-.0035	.465
	(16.3)	(26.7)	(13.9)	(-10.3)	(-9.7)	
Corrected						
	1.94	0.25	0.11	-.0012	-.0038	0.08
	(9.5)	(22.8)	(8.5)	(-5.9)	(-8.4)	(1.0)
Women						
OLS	2.49	0.21	0.07	-.0006	-.0028	
	(12.3)	(15.7)	(5.9)	(-3.2)	(-4.9)	.383
Corrected						
	1.19	0.27	0.12	-.0013	-.0043	.39
	(3.7)	(15.4)	(8.0)	(-5.9)	(-6.9)	(5.2)

t-ratios between brackets.

The results with and without the selectivity bias correction for the year 1990 are presented in Table 2[8] for illustrative purposes. As expected, correcting for the selectivity bias clearly appears to be quantitatively important, especially in the case of the female population. The coefficient of lambda, which reflects the relevance of the selectivity bias and the importance of the selectivity correction is always statistically significant and positive. The latter finding was an expected one because women with relatively lower human capital are typically those that

[8]Results do not differ for the other years.

are excluded from the sample. In other words, women with relatively low human capital have a lower propensity to participate in the labor market. Hence, the "inclusion" of these women tends to produce lower fitted wages, thereby increasing estimated gender wage gaps. The finding of a negative parameter associated with this variable, detected in other Latin American countries (see, for instance, Psacharopoulos and Tzannatos (1991)), has to be explained on the basis of the exclusion of educated women with relatively high levels of educational attainment from the labor market, a fact which leads to underestimating the fitted wage gaps[9].

The importance of the selectivity bias correction may be clearly highlighted with the 1990 results. Using the OLS results, the average observed male-female wage ratio was 1.023; that is, male workers received an hourly wage only 2.3 percent higher than female workers. However, as said before, the hourly wage sample (which considered the employed only) overstates the predicted population wage, since the employed labor force is mainly composed of "above average" females (or males).[10] However, this overestimation is smaller for males than for females. In fact, the corrected male-female wage ratio reached about 48 percent, instead of the estimated 2.3 percent. Based upon the corrected results, the sample wage for males was only 3 percent higher than the population wage, while the wage for females was 47 percent higher than the female population or corrected wage. Moreover, given that we excluded domestic services with a low human capital endowment from our sample, it is not surprising that the male-female wage ratio adjusted by human capital differences is 0.85. In other words, women in our sample have, on average, higher human capital endowments than men, so if men and women were paid according to the same non-discriminatory wage structure, the latter should earn 18% more than men. As working women actually earn considerably less than men, this evidence supports the idea that discrimination does in fact exist.

The selectivity bias correction indicates that, for the Chilean case, average and human capital corrected wages notably underestimate the discrimination effect against women. However, with regard to estimated human capital returns, quite a different story is portrayed. In the case of schooling, for example, rates of return estimated through uncorrected OLS generally favor males in about one point (e.g., in 1990 the uncorrected return to schooling for males and females was 17.6 percent and 16.5 percent, respectively). This situation changes dramatically when the wage equation is estimated through the selectivity bias correction. In this case, the female returns are consistently higher than those for males in all

[9] The Appendix offers an additional discussion on the findings in relation to the parameters.

[10] This conclusion is derived from the positive value of lambda and it means that observed workers are mainly those showing positive errors in the earnings equation (see the Appendix).

years considered.[11] This finding is consistent with the relatively lower schooling level prevailing in the case of the female population in LDCs. An important implication is that, contrary to common belief, human capital investment in females would be socially more profitable than in the case of males.

The female rate of return to experience increases relatively more than the return for males when the selectivity bias correction is introduced. Finding higher rates of return to experience in the case of females also has important policy implications. The effect of the higher discontinuity in the experience profile, expected in the case of females, would be considerably less important than what is usually believed.

TABLE 3
PREDICTED WAGE RATIOS

	Potential Experience	
	3 Years	15 Years
Male-Female Wage Ratio		
Predicted without selectivity bias correction	0.91	1.75
Predicted with selectivity bias correction	1.24	1.38

An hypothesis as to how discrimination against women is carried out would establish that it operates mainly through job promotions. According to this, the male-female wage ratio would notably increase with experience. To test this hypothesis we use the corrected and the uncorrected OLS regressions for 1990 and we analyze the wage differentials for two levels of experience evaluated at the average schooling. Seemingly supporting the mentioned hypothesis, observed uncorrected gender wage gaps increase very significantly with potential experience. This finding seems to support the idea that discrimination is mainly produced within the firm through a differentiated access to higher wage levels. However, when the selectivity bias correction is introduced, wage differentials are only slightly higher when experience increases (e.g., from, say, 3 to 15 years).

[11] In 1990, for example, the rate of return to schooling was 17.7% in the case of males, and 20% in the case of females. The complete series is presented in Table A1 (Appendix).

Therefore, the results using the corrected estimates of the return to experience do not support the idea that women are relatively excluded from the best jobs in their career. Instead, discrimination seems to be greater than it would appear from observed average wages, though it also does not seem to increase with the worker's experience.

4. WAGE DIFFERENTIALS OVER TIME

The likelihood of participating in the labor force, which is a function of human capital and labor market variables, is the critical factor that determines the existence of the selectivity bias. For instance, increasing participation rates in the case of the USA seem to explain why observed wage differentials have been falling so little over time, despite the government's efforts to reduce discrimination.[12] However, this interpretation cannot be directly extrapolated to LDCs, in particular to the Chilean case. In fact, in Chile LFP rates have notably increased throughout time, but it is not at all clear whether the greater female LFP corresponds basically to groups with relatively low labor productivity, as it is in the case of the USA. If, contrary to the latter experience, the increase in the rates of female labor force participation involved mostly "advantaged" women, the persistence in observed male-female wage gaps would be a clear symptom of increased discrimination.

The total observed (fitted) male-female wage gap, cannot be directly associated with discrimination. In the case of LDCs, in particular, it is clear that significant differences prevail both with regard to human capital stocks and rates of return to that human capital. Only the latter can be said to correspond to labor market discrimination. Hence, to measure discrimination, the Oaxaca's wage decomposition procedure was employed. Accordingly, the total male-female wage gap, obtained on the basis of the corrected OLS regression, was decomposed into a "pure discrimination" and a "human capital level" effects. The Becker's discrimination coefficient was used to measure the relative importance of the discrimination effect, and which is defined as:

$$D = [(Wm/Wf) - (Wm/Wf)^*] / (Wm/Wf)^* \qquad (3)$$

where Wm/Wf is the corrected male-female observed wage ratio --i.e., obtained on the basis of the Heckman-corrected OLS regression-- and the asterisk denotes the wage ratio estimated using women's average human capital, but paid according to the estimated rates of return for males.

[12] Smith (1989) analyzes the relative composition of employment in the USA, a variable which directly affects the magnitude of the sample selection bias.

Following Oaxaca (1973), the total observed wage differentials (G) and the differential associated with human capital endowments (N) can be measured as follows:

$$G = (Wm - Wf)/Wm \tag{4}$$
$$N = (Wm - Wf^*)/Wf^*$$

where the asterisk indicates that the female average human capital has been weighted by the measured rates of return for males. The above coefficients are related in the following way:

$$Ln(1+G) = Ln\ (1+D) + Ln(1+N) \tag{5}$$

This decomposition is applied for all the years used in this study (1958-1990) and used as well in the next section to analyze the time trend underlying the potential discrimination effect.

In assessing the effect of changing rates of labor force participation, a time series estimate of the discrimination coefficient is of capital importance. Consequently, a simple model was used to explain the behavior of the discrimination coefficient. This structure was analyzed on the basis of a series of observations over the period 1958-1990, and which totalled 31 observations across time of the discrimination coefficient, rates of return to human capital, labor force participation rates and human capital levels. The series of indicators for potential discrimination can be subsequently explained on the basis of an aggregate economic model.

Observed gender wage differentials have been declining notably over time. For instance, while in 1958 male wages exceeded female wages by 95 percent, in 1990 the corresponding percentage was only 47.4. This evolution may be explained by the reduction in human capital differences over time, as well as by the decline in market discrimination. In 1958 the "endowment effect," that is, the percentage of the observed wage gap explained by differences in human capital characteristics, favored males by 1.9 percent, whereas in 1990 the endowment effect favored females by 15.5 percent. In other words, the higher participation of women over time has been taking place mostly in terms of women with relatively higher human capital. On the other hand, the difference between the wage women would obtain if they were not discriminated against, and the wage they did actually obtain is 91.4 percent in 1958 and 74.4 percent in 1990. This suggests there has also prevailed an important reduction in the potential market discrimination, along with an equalization of human capital endowments.

The observed fluctuation throughout time of corrected gender wage gaps may be associated with i) the greater participation of women, especially those with relatively higher human capital, and ii) economic growth, which would produce tighter labor markets that make discrimination more expensive. Therefore, an interesting question, not as yet explored in the literature, is whether a consistent path in potential discrimination exists with regard to LFP and economic growth factors. This analysis could help to understand the origins of the market discrimination effect.

There is not an unique discrimination theory. What has been denominated "market" discrimination" is, in fact, an "unexplained" residual that is well worth exploring. Despite the lack of a comprehensive theory, two fundamental hypotheses can be tested against the data. The first hypothesis suggests that discrimination has to do with tastes (Becker, 1961). People would prefer men to women and hence firms would be willing to hire women only if there exists a (negative) productivity premium. The second hypothesis would establish that wage gaps mostly respond to cost differences, possibly associated with "protective female legislation." In Chile, for instance, a case in point would be those laws requiring firms to pay for child care and maternity leaves.

TABLE 4
FACTORS EXPLAINING THE MARKET GENDER DISCRIMINATION
(Dependent Variable: D)

	Constant	LISS	MWNC	GDPAdj.R2
Parameter	3.17	-.35	-.0036	.93.71
t test	(2.02)	(-2.29)	(-2.16)	(1.86)

D corresponds to the discrimination coefficient associated with the portion of measured wage gaps explained by differences in rates of return to human capital for males and females.

To test the relevance of these hypotheses, we estimated the effects of two variables on D over the 1958-1990 period. The first variable, the log of the market wage (LISS) attempts to capture the tightness of the labor market. A tight labor market is expected to increase the cost of discriminating against women, that produces a negative sign for the LISS coefficient. The second variable captures informality, that is, the degree to which firms comply with the law. This variable is measured by the ratio of minimum wages to the average wages of unskilled workers (MWNC). A larger value for this variable increases the probability of dealing with informal-unprotected workers. In that case, the presence of protective laws would be less relevant, and so would the need to pay wage compensations. Likewise, the GDP growth rate is also included in the model to test for the potentially important role played by cyclical economic fluctuations in relation to discrimination; for instance, periods of economic expansion should be associated with a higher degree of discrimination in terms of the lower cost which is, in turn, derived from the expanding activity, though keeping the real wages effect constant. The results of the OLS adjusted by first-order serial correlation are presented in table 4.

The empirical results indicate a clear pattern to explain the behavior of observed market gender discrimination. The existence of a tighter labor market helps to reduce the gender discrimination; furthermore, economic growth by itself will result in expanding gender gaps associated with pure discrimination. It is very likely, however, that economic growth will keep on reducing the male-female difference with regard to human capital stocks and hence contributing to reduce the total observed wage gap. Simultaneously, the degree of informality in the labor market, as measured by the ratio MWNC, will contribute to reduce the gender discrimination. The implication is that, if economic growth is associated with expanding formal labor markets relative to the informal segment, the gender discrimination will increase. This suggests that deregulating the labor market, as economic growth takes place, will be favorable to lower discrimination against women.

5. CONCLUSIONS

This paper has shown that to assess the extent of discrimination it is important to correct for selectivity in non random samples. There are three main findings. First, observed male-female wage differentials for a given human capital are underestimated in Chile because the sample of employed people excludes relatively less productive women. When this selectivity bias is empirically corrected, estimated male-female wage gaps are considerably larger. Second, human capital returns are not, as commonly believed (and observed), higher for males. On the contrary, selectivity bias corrections indicate that female returns are higher than those of their counterparts. Thus, more schooling would pay-off relatively better in social terms in the case of females. Finally, a consistent economic behavior of potential discrimination appeared across time. This path suggests that "protective" legislation has not necessarily been helpful for women, and that the most effective policy would be, instead, one that increases employment through neutral policies. In general, economic growth that arises in terms of higher wages and tighter labor markets, and that takes place in a less regulated environment, will have a more significant effect in terms of lower gender discrimination.

APPENDIX

The following equations, represent the market wage and the reservation wage respectively.

$$W_M = \alpha X + \mu_1 \qquad (1)$$

(1)

$$W_R = \beta X + \mu_2 \qquad (2)$$

where X is a human capital characteristic matrix, the α and β are vectors of parameters and the μ_i are "well behaved" random terms.

The researcher only observes wages when $W_M \geq W_R$. In terms of (1) and (2), this implies:

$$\alpha X - \beta Y > \mu_P \qquad (3)$$

where μ_p is $\mu_2 - \mu_1$ and is called "participation error". Normalizing this inequality by the standard deviation of the participation error we get:

$$I = \frac{\alpha X - \beta Y}{\sigma_P} > \frac{\mu_P}{\sigma_P} \qquad (4)$$

Given the selection criteria, the expected market wage is given by:

$$E(W_M / I > \mu_p/\sigma_p = \alpha X + E(\mu_1 / I > \mu_p/\sigma_p), \qquad (5)$$

and so it can be proved that,

$$E\,(W_M / I > \mu_p/\sigma_p) = \alpha X - \frac{\sigma_{1p}}{\sigma_P}\lambda \qquad (6)$$

where σ_{1p} is the covariance between μ_p and μ_1.

Given that we do not observe W_M independently of the value of I, Heckman has suggested to estimate (1) as:

$$W_M = \alpha X + \gamma\lambda + \mu \qquad (7)$$

where γ is the associated coefficient to λ $(-\sigma_{1P}/\sigma_P)$.

An important empirical question concerns the sign of the parameter γ. To this respect, consider that:

$$-\frac{\sigma_{1p}}{\sigma_P} = -\frac{Cov(\mu_{P1}\mu_1)}{\sigma_P} = -\frac{1}{\sigma_P} Cov\{(\mu_2-\mu_1),\mu_1\}$$

$$= -\frac{1}{\sigma_p} E\{\mu_2\mu_1-\mu_1^2\} = \frac{\sigma_1^2-\sigma_{12}}{\sigma_p}$$

$$= \frac{\sigma_1^2}{\sigma_P}(1-b\mu_2\cdot\mu_1) \qquad (8)$$

where $b\mu_2\ \mu_1$ is the coefficient from regressing μ_2 on μ_1. That is, given that $\sigma_1^2/\sigma_p > 0$, , the sign of γ depends on the magnitude of $b\ \mu_2\ \mu_1$, which is highly likely, be positive. This means that women with exceptionally high positive errors in the market wage (because they have a qualitatively better market oriented education for example), also have an exceptionally high reservation wage (because their very high "home productivity").

There is a slight possibility that this coefficient may have a value higher than one. In this case, the interpretation would be that women with exceptionally high positive market wages should show even higher reservation wages. Thus, the γ coefficient would be negative. However, there is no clear reason why this should be so. In the case of men, where "non observability" is mainly given by unemployment, we expect a small value for $b\ \mu_2\ \mu_1$ and hence a positive Lambda coefficient. A negative lambda coefficient would imply that non-participating women would be those with relatively low non observable errors. This idea, though reasonable in the case of some Central American countries, is not supported by the Chilean data that shows a positive correlation between highly paid professions with female participation, and also a positive correlation between family income and female participation.

TABLE A1

RATES OF RETURN TO SCHOOLING: UNCORRECTED AND CORRECTED MODELS

	Males				Females			
	Unco.	Corre.	Lambda	t test	Unco.	Corre.	Lambda	t test
1958	0.137	0.134	-0.085	-0.97	0.145	0.155	0.25	3.95
1960	0.121	0.122	0.029	0.36	0.138	0.145	0.18	2.70
1961	0.137	0.138	-0.170	-0.12	0.144	0.151	0.23	3.66
1962	0.138	0.138	0.054	0.68	0.143	0.155	0.39	5.89
1966	0.130	0.129	0.160	2.15	0.127	0.143	0.17	3.63
1967	0.121	0.133	0.083	0.98	0.138	0.142	0.20	3.86
1968	0.126	0.133	0.034	0.37	0.137	0.147	0.95	3.86
1969	0.139	0.139	0.046	0.51	0.136	0.148	0.25	4.12
1970	0.140	0.140	0.020	-0.26	0.147	0.165	0.31	5.30
1971	-0.014	0.149	0.080	1.21	0.115	0.162	0.21	4.10
1972	0.149	0.128	-0.020	-0.20	0.130	0.156	0.37	6.39
1973	0.117	0.118	0.050	0.71	0.112	0.126	0.25	4.84
1974	0.111	0.112	-0.180	-2.20	0.099	0.000	0.00	0.00
1975	0.111	0.112	-0.020	-0.24	0.110	0.130	0.24	4.18
1976	0.139	-0.490	-0.230	-2.19	0.130	0.161	0.18	3.49
1977	0.141	0.140	0.140	1.82	0.149	0.147	0.24	2.69
1978	0.145	0.150	0.070	0.90	0.126	0.156	0.33	5.39
1979	0.154	0.156	0.160	0.92	0.130	0.168	0.38	5.84
1980	0.157	0.156	0.080	0.98	0.137	0.165	0.31	4.85
1981	0.138	0.146	0.410	6.69	0.142	0.177	0.38	4.99
1982	0.156	0.189	0.580	9.91	0.150	0.195	0.45	5.35
1983	0.165	0.168	0.230	2.80	0.157	0.183	0.19	1.98
1984	0.172	0.000	0.000	0.00	0.160	0.185	0.00	0.00
1985	0.162	0.160	0.520	0.67	0.153	0.175	0.22	3.32
1986	0.161	0.169	0.510	0.70	0.160	0.182	0.24	3.49
1987	0.186	0.184	-0.130	-1.07	0.170	0.212	0.42	5.09
1988	0.172	0.173	0.260	3.30	0.162	0.145	-0.15	-1.70
1989	0.171	0.175	0.230	2.98	0.157	0.215	0.55	6.98
1990	0.176	0.177	0.080	1.04	0.165	0.200	0.39	5.16

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