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UNDERSTANDING INTERNATIONAL DIFFERENCES IN THE GENDER PAY GAP

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ABSTRACT

This paper tests the hypotheses that overall wage compression and low female supply relative to demand reduce a country's gender pay gap. Using micro-data for 22 countries over the 1985-94 period, we find that more compressed male wage structures and lower female net supply are both associated with a lower gender pay gap. Since it is likely that labor market institutions are responsible for an important portion of international differences in wage inequality, the inverse relationship between the gender pay gap and male wage inequality suggests that wage-setting mechanisms, such as encompassing collective bargaining agreements, that provide for relatively high wage floors raise the relative pay of women, who tend to be at the bottom of the wage distribution. Consistent with this view, we find that the extent of collective bargaining coverage in each country is significantly negatively associated with its gender pay gap. Moreover, the effect of pay structures on the gender pay gap is quantitatively very important: a large part of the difference in the gender differential between high gap and low gap countries is explained by the differences across these countries in overall wage structure, with another potentially important segment due to differences in female net supply.

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I. Introduction

Virtually every industrialized country has passed laws mandating equal treatment of women in the labor market. Yet the gender wage gap, while on the decline in many countries, is a persistent feature of virtually every nation's labor market. Moreover, the extent to which men outearn women varies substantially across countries as well. Among the sample of 22 countries shown in Table 1 below, the gap in log earnings corrected for differences in weekly work hours between men and women averaged .307 log points over the 1985-94 period.[1] The standard deviation of these gaps across countries was .145 log points, and the interquartile range (75-25 difference) in gender pay gaps across countries was .124 log points. In this paper, we attempt to explain these international differences.

Economists have traditionally looked to gender-specific factors such as female shortfalls in human capital or employer discrimination against women to explain the size of the gender pay gap and its evolution over time. However, beginning with the work of Juhn, Murphy and Pierce (1991), economists have recognized that overall wage structure, or the prices the labor market attaches to skills and the rents accruing to those in favored sectors, can have a major impact on the relative wages of different subgroups in the labor market. The logic behind this insight is straightforward. For example, since women have on average less labor market experience than men and tend to work in different occupations and industries, an increase in the return to experience or in sectoral differentials will raise the gender pay gap, all else equal. As we have pointed out in our earlier work (Blau and Kahn 1992; 1995; and 1996b), the same reasoning applies across countries: countries with relatively high rewards to skill and relatively large sectoral differentials will tend to have larger gender pay gaps, ceteris paribus. International comparisons provide a particularly fertile field in which to study the effects of wage structure on pay differentials between men and women because, due in part to substantial international differences in wage setting institutions (Blau and Kahn 1996a), they provide greater variation in

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[1] This is an unweighted average of the raw gender pay gaps by country presented in Table 2. Below, we explain in detail how the estimated pay gaps were obtained.

wage dispersion and labor market rewards than may generally be obtained by shifts in wage structure over time within a country. An important implication of this approach is that the factors that influence overall wage structure such as supply of and demand for skills and wage-setting institutions may well be extremely important determinants of international differences in the gender pay gap and, by implication, of trends in the gender pay gap within countries as well.

In previous work on international differences in the gender pay gap (Blau and Kahn, 1992, 1995; 1996b), we used this insight to address a paradox: while the relative qualifications of American women are high compared to women in other countries and the United States has had a longer and often stronger commitment to anti-discrimination laws than most economically-advanced nations, the United States has traditionally been among the countries with the largest gender gaps. Due in part to its decentralized wage setting institutions, the US labor market places a much larger penalty on those with lower levels of labor market skills (both measured and unmeasured) or who are located in less-favored sectors of the labor market. These are disproportionately women, whose relative wages are thus potentially reduced by labor market institutions in the US. Based on pair-wise comparisons of the United States to nine other advanced countries, we presented evidence that the higher level of wage inequality in the United States is the primary reason for its relatively high gender pay gap.[2]

Specifically, using a decomposition technique developed by Juhn, Murphy and Pierce (1991), we simulated the gender pay gap that would result if other countries had the overall wage structure of the US. We found that each simulated "other country" gender pay gap would be virtually as large as or larger than the US gap under this counterfactual, implying that differences in overall wage structure usually completely accounted (or in several cases more than completely accounted) for the US-other country differences in the gender pay gap.[3] In these comparisons, it

[2] The countries were Australia, Austria, Britain, Hungary, Italy, Norway, Sweden, Switzerland, and West Germany.
[3] This framework has been subsequently adopted by other researchers who have also found evidence of the impact of wage-setting institutions on the gender gap in pay. Kidd and Shannon (1996) found that 37-66 percent of the smaller gender pay gap in Australia than Canada in 1989-90 (a difference of .14 log points) was due to Australia’s more compressed wage structure. Australia’s institution of nationally-binding wage awards issued by government tribunals as well as its higher level of unionization are likely candidates for explaining its lower wage dispersion. Similarly, Edin and Richardson (1999) found that when Sweden's solidarity bargaining was compressing the wage

was clear that relying on a framework that only considered gender-specific factors would have been inadequate to explain the relatively high gender pay gap in the United States.

In this paper, we extend this analysis asking whether wage structure has an important effect on the gender pay gap across a broad variety of countries. That is, while wage structure may be the most important reason for the relatively high gender pay gap in the United States, does this framework explain a large component of international differences in the gender pay gap generally? Moreover, while our earlier research using two-country comparisons and the Juhn, Murphy, and Pierce (1991) decomposition suggested a strong role for wage inequality and wage-setting institutions in affecting the gender gap, it did not directly test this effect. The technique assumes that women will be affected by the same forces that influence the male wage distribution in a country. Specifically, the estimated male prices of measured characteristics are assumed to affect men and women in the same way and the residuals from a male wage regression are decomposed into a portion reflecting the prices of unmeasured skills and a portion reflecting the quantities of unmeasured skills. Taking the US and Sweden as an example, the decomposition of the residual assumes that, controlling for measured characteristics, if Swedish women with wage residuals equal to, say, the 35th percentile of the Swedish male residual wage distribution were moved to the US they would have wage residuals equal to the 35th percentile of the American male residual wage distribution. This type of decomposition has been criticized by Suen (1997). In this paper, we are able to implement a more direct test of this idea which provides considerable empirical support for our earlier findings and thus indirectly for the Juhn, Murphy, and Pierce decomposition technique.

Some support for the notion that wage structure is important may be found in Rowthorn's observation, based on tabulations of aggregate wage data on 17 countries in 1973 and 1985, that the gender pay gap tended to be lower in countries with more centralized wage setting

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distribution at the bottom during the 1968-74 period, the gender pay gap in Sweden fell by .06 log points; 82% of this decline was due to the compression of the wage structure.

institutions as measured by industrial relations researchers (Calmfors and Driffill, 1988).[4] In addition, in earlier work, we found a positive (though insignificant) correlation between a country's gender pay gap and the degree of decentralization of its wage setting across a small sample of ten countries primarily from the late 1980s (Blau and Kahn, 1996b).

In this study, we use micro-data on 22 countries over the 1985-94 period to test whether more egalitarian pay structures (as measured by the extent of male wage inequality) are associated with lower gender pay gaps. In contrast to studies based on published data from sources such as the ILO or the OECD, our micro-data on individual workers allow us to adjust for international differences in women's human capital in estimating the gender pay gap, an important step in isolating the impact of wage structure. Data on individuals are also essential in estimating the characteristics of the wage structure itself if we are to separate out labor force heterogeneity from the true effects of labor market prices. A further contribution of the current study is that we are able to estimate the impact of supply and demand conditions in the female labor market on a country's gender pay gap.

Our results provide support for the idea that egalitarian wage structures reduce the gender pay gap. We also find that, in some models, controlling for wage structure, the gender differential is lower when women are in shorter supply relative to the favorableness of the country's demand structure for women (female "net supply"). These results hold equally strongly even when we exclude the United States from the analysis, suggesting that the negative relationship between wage compression and the gender pay gap is not driven by the extreme case of the US. Thus, to the extent that institutions are a major factor leading to more or less egalitarian wage structures (Blau and Kahn, 1996a), we find strong evidence for the importance of institutions and some evidence for the effect of market forces in affecting the gender pay gap. Moreover, the effect of pay structures on the gender pay gap is quantitatively very important: a

[4] Rowthorn (1992) did not perform a statistical test of this hypothesis. However, using his reported data, we regressed a country's female/male hourly earnings ratio on its centralization rank (1=most centralized) and found a coefficient of -.931 with a standard error of .410 in 1985; and a coefficient of -.946 with a standard error of .307 in 1973. Thus there appears to be a stable negative relationship at the aggregate level between decentralization of wage-setting and the female/male pay ratio in the 1970s and 1980s.

large part of the difference in the gender differential between high gap and low gap countries is explained by the differences across these countries in overall wage structure, with another potentially important segment due to differences in female net supply.

II. The Institutional Setting

In this section we briefly summarize international differences in gender-specific policies and basic wage-setting institutions and their expected effects on the gender pay gap. Human capital is also a major determinant of gender pay gaps, and, in the work below, we control for international differences in women's measured human capital compared to men. However, international variation in policies and institutions appear to be more dramatic than those in women's relative human capital levels, at least in our sample. Further, human capital can be affected by such policies and institutions as discussed below. Differences across countries in institutions which affect the gender pay gap may be classified into those that are gender-specific and those that affect the wage structure in general.

Gender-specific policies include equal employment opportunity (EEO) and anti-discrimination laws, as well as laws and policies governing parental leave and child care availability. The expected positive effect of EEO policies on the gender earnings ratio is reasonably straightforward, although the impact will most likely depend on the effectiveness of the legislation's enforcement as well as its provisions. In general, it is expected that, given considerable segregation of women by occupation, firm, and industry, equal pay laws mandating equal pay for equal work within the same occupation and firm will have a relatively small effect. Laws requiring equal opportunity, hiring preferences, and/or "comparable worth" (i.e., equal pay for work of equal value to the firm, regardless of specific occupational category) have potentially larger impacts. In addition, since EEO laws involve occupational shifts, they may require considerable time to have an impact on pay. Thus, the comparable worth approach which provides for immediate increases in relative pay in female-dominated occupations may be expected to have the largest initial wage effect, possibly accompanied by a negative impact on female employment.[5]

Virtually all OECD and European Community countries had passed equal pay and equal opportunity laws by the mid-1980s, although the US implemented its anti-discrimination legislation before most other countries (OECD, 1988, pp. 167-168; Simona 1985). Of the Eastern European countries in our sample, only Russia and Hungary were listed by the ILO (1994) as having anti-discrimination laws as of 1994; the other Eastern European nations in our sample—Bulgaria, Czech Republic, Poland, and Slovenia—were not included in the ILO listing.[6] However, Brainerd (2000) reports that under socialism in Eastern Europe and the Soviet Union, there were government guarantees of equal pay for equal work, and women's labor force participation was encouraged. One country with perhaps the strongest intervention against sex discrimination is Australia, the only one to have implemented a national policy of comparable worth through its labor courts (Gregory and Daly, 1991; Killingsworth, 1990).[7] Comparable worth pay policies remain rare in the US private sector, although they have been adopted by a number of state and local governments (Blau, Ferber and Winkler 1998).

The expected impact of family leave (disproportionately taken by women even when it is available to men) is unclear *a priori*. On the one hand, such policies may raise the relative earnings of women by encouraging the preservation of their ties to particular firms and hence increasing the incentives of employers and women workers to invest in firm-specific training. On the other hand, the existence of such policies could increase the incidence and/or duration of temporary labor force withdrawals among women, raising the gender gap for the affected group. Further, the incremental costs associated with mandated leave policies may increase the

[5] For evidence on these employment effects, see, for example, Killingsworth (1990) and Gregory and Duncan (1981).
[6] Russia's law prohibited discrimination against pregnant women.
[7] Switzerland incorporated the principle of equal pay for work of equal value into its constitution in 1981 (Simona 1985), but there is no indication that it has been implemented as yet. And, in 1991, Ontario required virtually all employers to practice the principles of comparable worth, but other Canadian provinces have not passed such legislation (Gunderson and Robb 1991).

incentives of employers to discriminate against women. In Germany, for example, where mandated parental leave is of long duration, it was reportedly legal for employers to discriminate against pregnant women (Demleitner 1992).

By international standards, the United States has a relatively weak entitlement to parental leave, consisting of an unpaid 13-week mandated period, which was only introduced in 1993. In contrast, other OECD countries implemented mandated leave earlier and most have a much longer period of leave, usually paid (Ruhm, 1998). While the theoretical effect of parental leave on the gender gap is ambiguous, some research on the impact of parental leave has found a positive effect of short leave entitlements on women's relative wages, although extended leaves have been found to have the opposite effect (Ruhm, 1998; Waldfogel, 1998). In some specifications below, we are able to control for the duration of parental leave entitlement, which in our data ranged from 0 to 156 weeks, although these data are not available for our full set of countries.

Child care is another important area of public policy which particularly affects women, but one which is difficult to summarize across a large set of countries. Although we were unable to secure data for an extended set of countries, some available evidence suggests that, as of the mid-1980s, the United States had a smaller share of young children in publicly funded child care than many other OECD countries, but provided relatively generous tax relief for child care expenses (Gornick, Myers and Ross, 1997).

With respect to wage structure, there is a considerable body of evidence showing that wage-setting institutions affect a country's level of wage inequality. These institutions take the form of collective bargaining conventions, minimum wage laws, and governmentally mandated extensions of the terms of collective bargaining agreements to nonunion workers (Blau and Kahn 1999). In general, more heavily unionized economies in which collective bargaining takes place at highly centralized levels have been found to have the lowest overall wage dispersion (Freeman 1988; Rowthorn 1992; Blau and Kahn 1996a).

It seems likely that systems of centrally-determined pay also entail smaller gender wage differentials for at least two reasons related to their impact on skill prices and sectoral differentials. First, in the US, a significant portion of the male-female pay gap is associated with interindustry or interfirm wage differentials that result from its relatively decentralized-pay setting institutions (Blau, 1977; Johnson and Solon, 1986; Sorensen, 1990; Groshen, 1991; and Bayard, Hellerstein, Neumark, and Troske, 1999). Thus, centralized systems which reduce the extent of wage variation across industries and firms are likely to lower the gender differential, all else equal. Second, since in all countries the female wage distribution lies below the male distribution, centralized systems that consciously raise minimum pay levels, regardless of gender, will also tend to lower male-female wage differentials. Finally, the impact of gender-specific policies to raise female wages may be greater under centralized systems where such policies can be more speedily and effectively implemented.[8]

There is considerable variation across countries, both in collective bargaining coverage and in the degree of centralization of wage-setting in both the union and nonunion sectors. Among the OECD nations, the Scandinavian countries and Austria stand at one extreme with their high degree of collective bargaining coverage and union-negotiated wage agreements at the economy-wide or industry-wide level, while the US stands at the other extreme with an especially low rate of collective bargaining coverage and pay setting which is often determined at the plant-level even within the union sector. Moreover, in many of the OECD countries, but not in the US, formal or informal mechanisms exist to extend union-negotiated pay rates to nonunion workers. In some specifications below we look directly at the effect of collective bargaining coverage on the gender pay gap and find evidence of the expected negative effect.

Systems of pay compression may influence workers' incentives to invest in skill acquisition, although the direction of these effects is theoretically ambiguous. On the one hand, if the return to experience is reduced by wage compression, women will have less incentive to

[8] An example of this is provided by the case of Australia when the adoption of comparable worth by the labor courts produced the rapid increase in the gender earnings ratio (e.g., Gregory and Daly, 1991; Killingsworth, 1990).

remain employed during periods of childbearing and childrearing than in countries where there is a larger penalty for time out of the paid labor force. In our data, as well as in most microdata, we are unable to measure actual experience; thus, this type of labor supply response of women to pay compression will go largely undetected. This could bias us away from finding a positive effect of reduced male wage inequality and wage-compressing institutions on the relative pay of women, controlling for their measured, as opposed to actual, human capital levels. On the other hand, high wage floors may encourage women's labor force participation, raising their relative experience levels. Of course, these wage floors could still reduce women's employment if labor demand effects are important. In any case, wage-induced changes in female qualifications could produce a positive or a negative bias on our estimates of the impact of wage compression.

Another institution that directly affects the wage distribution is mandated minimum wage coverage. Studies of the impact of minimum wages in various countries on the wage distribution invariably find that such regulation compresses the bottom of the distribution (Card and Krueger 1995; DiNardo, Fortin and Lemieux 1996; Machin and Manning 1994; Katz, Loveman and Blanchflower 1995; Dolado, et. al 1996). And Blau and Kahn (1997) concluded that falling real minimum wages over the 1980s retarded the narrowing of the pay gap between low-skill women and low skill men. International differences in minimum wage levels, both those negotiated in collective bargaining and those imposed by government, are likely to be reflected in our measures of male wage inequality. Below, we also present some results designed to explicitly assess the impact of minimum wages on the gender pay gap.

An alternative perspective on these issues is provided in a recent paper by Fortin and Lemieux (1998). They suggest that relative gains for women can affect male wage inequality, in effect posing the opposite direction of causality to the one posited here.[9] As they point out, their approach is consistent with a pure job assignment model in which alterations in the relative position of women in the wage distribution do not affect the overall distribution of wages. This

[9] We note that this reasoning applies whether women's gains come through reductions in discimination or through improvements in women's relative human capital levels.

raises the possibility that a reduction in the gender gap could increase male inequality (as some women move up the wage distribution and some men move down); and they present some evidence consistent with this for the US over the 1980s. To the extent that female gains do come to some degree from male losses, this reverse causation will impose a negative bias on the measured effect of male inequality on the gender gap. In addition, in some specifications presented below, we attempt to avoid the simultaneity problem by taking a reduced form approach in which institutional variables such as collective bargaining coverage, minimum wages, unemployment insurance generosity, and the degree of employment protection take the place of our measures of male inequality and female net supply.

III. Description of the Data

In this study, we use microdata from the 1985-94 annual files of the International Social Survey Programme (ISSP). The ISSP is a voluntary grouping of study teams in over twenty countries "each of which undertakes to run a short, annual self-completion survey containing an agreed set of questions asked of a probability-based, nation-wide sample of adults." (ISSP documentation file, 1992). In addition to its annual theme questions concerning social attitudes, each year the ISSP administers a uniform set of questions on respondents' education, earnings, age, marital status, work hours and gender. The sample design is a new random cross-section in each available year for every country, with the exception of data for Australia in 1986, 1987, 1991, and 1994, and Hungary in 1990-94 which are panel follow-up samples. Our results were virtually identical when we excluded these country-year observations from the data.

The sample of countries with available wage data varies from year to year and is shown in Table 1.[10] Each country appears an average of 4.5 times during this period yielding a total of 100 country-year observations. As may be seen in the table, the coverage of countries has grown

[10] Data were also available for Spain in 1993, but were not included here because the estimated gender gap, .061, was not credible in comparison to other sources. Data on Australia for 1985 were also excluded due to data problems noted in Blau and Kahn (1995).

from a relatively small number (four) in 1985 to seventeen countries with usable data in 1994. A particularly interesting feature of the ISSP sample is the addition of several Eastern European countries, including the former East Germany, the Czech Republic, Hungary, Bulgaria, Russia, Slovenia, and Poland, mostly after 1990.

Table 1 also shows the earnings concept for each country. Unfortunately there are some differences across countries in the exact definitions of the income variables in the ISSP data. While in over half the cases the income data are measured before taxes, income is measured after taxes in the remainder. Although this difference is of some concern, it is reassuring that the empirical results reported below were not affected by the inclusion of a dummy variable indicating that net income was the income concept employed, nor was the coefficient for this dummy variable ever significant. In addition, because of the different possible tax treatment of single vs. married women and the issue of before vs. after tax earnings, we estimate some supplementary models confining ourselves to married workers only, in addition to our basic specification that includes all workers. Our results are very similar for the married subsample, providing some further indication that differing income definitions do not have a major impact on our results.

As may also be seen in the table, with the exception of Ireland in 1993 and 1994, earnings are computed on either an annual or a monthly basis; thus, in general, we are not able to compute hourly or weekly earnings (the ISSP does not collect information on weeks worked). However, we do have information on hours worked per week, which as described below, we use to estimate earnings adjusted for time input. Since the national surveys that comprise the ISSP can change their focus from year to year, variables such as industry and union membership are not consistently available. However, occupational information is available for most country-year observations and we use it in several of our analyses. Finally, it may be noted that not every country participates in every year, a factor that further contributes to the unbalanced nature of the country-year panel.

IV. Estimation Strategy

Our analysis begins with log earnings equations estimated separately by sex for each country j and year t, where i indexes individuals:

$$(1)\qquad LnEARN_{ijt} = b_{0jt} + b_{1jt}PART_{ijt} + b_{2jt}HRPART_{ijt} + b_{3jt}HRFULL_{ijt} + X_{ijt}B_{jt} + e_{ijt},$$

where LnEARN is the natural log of earnings; PART is a dummy variable for part-time employment defined as less than 35 hours per week;[11] HRPART and HRFULL are interactions of weekly work hours with part- and full-time status; X is a vector of explanatory variables described below; and e is an error term. The model allows for both a part-time shift term and different slopes for hours for part-time and full-time workers. A detailed adjustment for part-time employment is important in light of the prevalence of part-time work for women in many countries. We use a linear specification for the hours variable and its interactions with the part time and full time dummies as part of a relatively unrestricted simple specification intended to capture the variation in labor input, given that we do not have information on weeks worked. Results were virtually identical when we used log hours instead of hours in equation (1).

The PART, HRPART and HRFULL coefficients from (1) were used to adjust each person's earnings for work hours by assuming a 40 hour work week. That is, for each worker i, we have (suppressing the country and year subscripts):

$$(2)\qquad LnEARN(40)_i = lnEARN_i - b_1PART_i - b_2HRPART_i - b_3(HRFULL_i - 40)$$

where the coefficients, b_n, are obtained from estimating equation (1) for males (m) and females

[11] Countries may differ in their definitions of part time. Until 1997, the OECD used each country's own definition in reporting on the fraction of workers employed part-time. For our data period (1985-1994), this was 35 hours in most cases, leading us to use the 35 hour definition here. Starting in 1997, in recognition of some recent trends in the definition of full-time work, the OECD began using a common 30 hour cutoff in its statistics (see, for example, OECD 1999). Because of changing norms with respect to what is considered full time work, we tested the robustness of our results by using a 30 hour cutoff and obtained very similar findings to those reported below.

(f) separately. Thus, the gender difference in the mean of LnEARN(40) in a given year and country is our best estimate of the gender gap in wage rates.[12]

The explanatory variables in X include the traditional human capital variables of education and potential experience and its square, as well as a constant. Inclusion of marital status as a human capital characteristic is problematic in that it is likely to be positively associated with productivity for men but may be negatively so for women (e.g., Korenman and Neumark, 1991). Because of this ambiguity, we do not control for marital status in the country-year microdata wage regressions used to estimate human capital-adjusted gender pay gaps; however, similar results were obtained when we did control for marital status and, as noted above, when we confined the sample to married individuals. In our main specification, both the self-employed and wage and salary workers are included in the regression sample, but, as discussed below, results were similar when the sample was restricted to wage and salary workers only. The rationale for including the self-employed is that the size of this sector may be influenced by wage setting institutions. For example, in economies where wage determination is heavily regulated, the self-employment sector may swell in an attempt to avoid such regulation.

In additional specifications for a subset of the country-year observations, we are able to estimate models that control for one-digit occupation.[13] We recognize that an individual's occupation is potentially endogenous because individuals may choose jobs on the basis of the opportunities available; further, a country's occupational structure may be affected by its wage structure. However, occupation is also likely to be an important indicator of human capital and may serve as a proxy for actual experience, which is not available in the ISSP. Thus, while we present results not controlling for occupation, we also estimate supplementary specifications with occupational controls. Our basic conclusions are robust with respect to this addition.

Having estimated (1) and (2) for each country and year (100 regressions for men and 100

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[12] We used this methodology in our earlier work (Blau and Kahn, 1992; 1995; 1996b) and obtained estimated gender pay gaps that were similar in country rankings to those reported in published data based on average hourly earnings.

[13] As noted above, data on industry and unionism are not consistently available across our sample of countries.

for women), we then construct three different measures of the gender gap in hours-corrected earnings. We begin with the estimated raw gender log wage gap:

$$(3) \qquad \text{TOTAL GAP}_{jt} \equiv \text{LnEARN}(40)_{mjt} - \text{LnEARN}(40)_{fjt},$$

where $\text{LnEARN}(40)_{mjt}$ and $\text{LnEARN}(40)_{fjt}$ are, respectively, the average male and female levels of log full-time earnings for country j and time t. The raw gender gap represents the net effect of all of the forces that influence the gender gap in hours-corrected earnings. These include gender-specific effects influencing the male and female values of the explanatory variables X, the impact of discrimination, and the impact of labor market prices as discussed above. Further, by not conditioning on the X variables, when we use the raw gender gap as the dependent variable, we allow the country's labor market prices to affect men's and women's incentives to acquire these measured skills. As we have seen, wage compression may have opposing effects on women's relative qualifications. Using TOTAL GAP as the dependent variable allows the net effects of these processes to be observed.

While wage compression may, as argued above, affect women's and men's relative values of X, there may also be exogenous (e.g., cultural) reasons for men and women to have different relative levels of qualifications in different countries. For this reason, we primarily focus on results based on the GAP US CHARS which estimates the predicted gender pay gap on the assumption that the men and women in each country-year microdata file have the same average levels of measured qualifications as US men and women for that year:

$$(4) \qquad \text{GAP US CHARS}_{jt} = (40b_{3mjt} + X_{mut}B_{mjt}) - (40b_{3fjt} + X_{fut}B_{fjt}),$$

where u refers to the United States, X is a vector of means of the explanatory variables for the indicated group, and the subscripts f, m, and j have been defined above. Equation (4) provides a simulated gender pay gap that removes international differences in women's relative levels of measured characteristics. However, since gender differences in measured characteristics remain (i.e., at the US levels) there is ample room for each country's wage structure to influence the magnitude of the gender pay gap through its effects on the returns to the measured Xs. Finally, note that equation (4) simulates full time earnings by valuing the country's own male and female full-time hours coefficients b_{3mjt} and b_{3fjt} at 40 hours.

A second specification of the gender gap that standardizes for international differences in women's relative qualifications is the familiar "unexplained" gender pay gap:

$$(5) \qquad \text{UNEXPLAINED GAP}_{jt} = (40b_{3mjt} + X_{fjt}B_{mjt}) - (40b_{3fjt} + X_{fjt}B_{fjt}).$$

Equation (5) tells us the size of the gender pay gap in country j in year t that results from unequal rewards for men and women of equal measured characteristics, using the country's female mean characteristics as weights. This is often taken to be a measure of labor market discrimination, although it will be influenced by gender differences in unmeasured characteristics as well. Differences in the degree of wage compression across countries may influence the size of the UNEXPLAINED GAP in at least two ways. First and most straightforwardly, to the extent that there are gender differences in unmeasured characteristics the size of the UNEXPLAINED GAP may be affected by international differences in the prices of these unmeasured factors. Second, if discrimination takes the form of treating women of a given skill level as being comparable to men with a lower skill level, compression of wages will reduce the wage consequences of such actions. Although we examine some results for the UNEXPLAINED GAP, for the reasons detailed above, we believe the GAP US CHARS is the more appropriate measure.

Having described the construction of the basic dependent variables, we turn to the major explanatory variables: male wage compression and the favorableness of supply relative to demand for female labor. With respect to male wage compression, the simplest approach would be to use a measure of observed male wage inequality. However, observed male wage inequality is influenced by heterogeneity of productivity characteristics as well as by labor market prices, and it is prices that generate the compression of wages that are associated with collective bargaining institutions. To focus on prices, we construct two measures of male wage inequality that remove the effects of measured heterogeneity.[14] First, for each country and year, we take the US sample of men for that year and compute a predicted log wage for each US male (ui) using the male price structure for country j and year t:

(6) $$(X_{us})_{ijt} = 40b_{3mjt} + X_{uit}B_{mjt}.$$

We then compute measures of inequality, including the standard deviation and the 50-10 gap across the American male sample (the ui's) of these $(X_{us})_{ijt}$, which gives us a measure of the degree of male inequality which results from the prices of measured human capital in country j in year t. In this case we include a control for marital status in X since there is less ambiguity about the effect of this variable among men. The rationale behind examining the 50-10 gap is that it is a measure that may be especially relevant to those at the bottom of the distribution, such as women. Note that we use the mean of $(X_{us})_{ijt}$ to compute the male part of the GAP US CHARS in equation (4). We use the inequality in the distribution of the $(X_{us})_{ijt}$ as an indicator of the compression caused by human capital prices.

Second, an insight of Juhn, Murphy and Pierce's (1991) work is that wage residuals also potentially contain information about labor market prices, in this case the prices of unmeasured characteristics. Accordingly, we also include as an explanatory variable the standard deviation of (or the 50-10 gap in) the residuals from the male log wage equation estimated for country j in year t. Note that the regression residual includes not only the returns to unmeasured individual characteristics, but also industry and firm wage effects which we are not able to capture due to data limitations in the ISSP. There is evidence that industry wage effects vary significantly

[14] We have also estimated our models with the raw 50-10 male wage gap as the indicator of wage compression with very similar results to those reported below.

across countries (Edin and Zetterberg 1992; Barth and Zweimüller 1992; Kahn 1998) and firm wage effects are believed to do so as well (Teulings and Hartog, 1998).[15] However, in addition to containing information about unmeasured prices, residual inequality is also affected by other factors including heterogeneity in unmeasured characteristics and measurement errors, while the inequality in predicted wages is a direct measure of labor market prices.

In addition to labor market prices and relative human capital levels, we also expect supply and demand conditions for women relative to men to influence the gender pay gap. If men and women are viewed as imperfect substitutes in the labor market, as assumed in supply and demand models of wage inequality (e.g., Katz and Murphy, 1992; Juhn, Murphy and Pierce, 1993; Bound and Johnson, 1992), then countries with higher demand for, or lower supplies of, women relative to men will have smaller gender pay gaps, other things equal. To measure these demand and supply factors, we construct demand and supply indexes for women relative to men in a manner similar to those employed in the wage inequality literature.

Turning first to demand, we wish to know, for each country-year observation, how favorable the composition of output by industry (and the consequent derived demand for labor) is for women. Following Katz and Murphy (1992), we constructed the demand index, $\ln(1+\Delta D_{jt})$, for women for each country j and year t relative to a base of 1989 in the US:

$$(7) \qquad \Delta D_{jt} = \Sigma_i c_{iu89}(E_{ijt} - E_{iu89})/E_{iu89},$$

where i refers to one digit industry cell, c_{iu89} is the fraction of all US women workers who were employed in industry i in 1989, E_{ijt} is the share of total employment in industry i, for country j in year t, E_{iu89} is share of total employment in industry i in the US in 1989.[16] The demand index

[15] The increasing availability of matched firm-worker data should facilitate future efforts to assess international differences in the magnitude of these firm effects (Abowd and Kramarz 1999).
[16] Since data on industry were not consistently available across countries in the ISSP, we used published sources to obtain this information for each country and year; see ILO (various issues).

thus measures the degree to which the country j year t industry structure favors women relative to the US in 1989, using US weights.

Supply indexes ΔS_{jt} are computed as follows:

(8) $$\Delta S_{jt} = \ln F_{jt} - \ln F_{u89},$$

where F_{jt} is the share of country j's workers in year t who were women, and F_{u89} is the analogous fraction for the US in 1989. Thus, the supply index shows the relative representation of women in country j's work force, using the 1989 US share as the norm.

We may then compute net supply as:

(9) $$\Delta NS_{jt} = \Delta S_{jt} - \ln(1+\Delta D_{jt})$$

where ΔD_{jt} and ΔS_{jt} are defined in equations (7) and (8) above. (Recall that all magnitudes are in log points and have been normalized relative to the US in 1989.) As implied by Katz and Murphy's (1992) equilibrium model, differences across countries in the gender pay gap will be positively related to differences in female net supply, ΔNS_{jt}. Intuitively, the larger the supply of women relative to demand in country j compared to the US, the worse women will fare in country j compared to the US.

While this net supply index represents our best measure of the relative abundance of women workers, we note that it is overstated (understated) for countries with relatively small (large) gender pay gaps. This is the case because, while we seek a measure of the position of the female relative supply curve in relation to the relative demand curve for female labor, our measures are based on employment outcomes. This means that countries with a low (high) gender pay gap will appear to have larger (smaller) female supply and smaller (larger) female demand indexes than otherwise because of the movements of workers and firms along their supply and demand schedules. Thus, the construction of the net supply index may lead to a bias away from finding that relative scarcity of women raises their relative wages. Another problem with this measure is that it does not reflect differences across countries or over time in the within industry demand for women workers. Nonetheless, it is important to control for relative net supply, since this is potentially a key determinant of women's relative wages. Further, Topel (1994) has argued that rising female labor supply has in fact contributed to rising male wage inequality. This is based on the idea that women compete in the labor market with low skill men.[17] To the extent that this argument holds, our control for female net supply provides a sharper test of the effect of male prices on the gender pay gap. We interpret the impact of male prices as being their effect beyond any portion that may have been caused by female supply.

A final reason for including a measure of net supply concerns possible selection bias due to the fact that we observe earnings only for those with jobs. In addition to adjusting for the effect of relative scarcity on the gender gap, our control for net supply also accounts for the impact of the selectivity of the female and male labor forces on the gender pay differential. Note that we do not impose any functional form on this relationship, or even posit its sign. While it is most straightforward to expect those in the labor market to have better wage offers than the nonemployed, this need not be the case. Especially for women, it is possible that those out of the labor force have better labor market opportunities than those at work, on average, but their home productivity may be so high as to outweigh their higher wage offers. If this is the case, selectivity will impart a positive bias to the estimated effect of female net supply on female relative wages. Conversely, if the nonemployed have worse wage offers, then selectivity will bias us toward a negative relationship between female net supply and female relative wages. While our control for net supply makes us more confident than otherwise that selectivity biases are not driving the results for the other explanatory variables, consideration of this issue implies

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[17] In earlier international work (Blau and Kahn 1996a), we did not find evidence of a consistent relationship between female net supply and male wage inequality. Similarly, Juhn and Kim (1999) found little evidence that female supply affected male wage inequality within the United States; see also Blau and Kahn (1997).

that we must be cautious about concluding that supply and demand have important effects on the gender pay gap.

Labor market prices and female net supply form the basic explanatory variables for our study of the determinants of the gender pay gap, as expressed in estimating equations of the following form:

$$(10)\quad \text{GENDER GAP}_{jt} = c_0 + c_1\text{INEQ }(X_{us})_{jt} + c_2\text{INEQ RESID}_{jt} + c_3\text{NETSUPPLY}_{jt} + \Sigma_s d_s \text{YR}_{st} + e_{jt},$$

where for each country j and year t, GENDER GAP is a measure of the gender log wage gap; INEQ (X_{us}) is male inequality in measured prices; INEQ RESID male residual inequality; NETSUPPLY is female net supply (as defined in equation 9); YR_s is a dummy variable for year s, with a range of 1986 to 1994 (1985 is the omitted year); e is an error term.[18]

To test for the robustness of our findings with respect to wage compression and net supply, we add a number of controls for other policies, cultural differences, or other gender-specific factors that may influence the male-female pay gap. First, we control for whether a country was in the former Eastern bloc and test whether the impact of compression and net supply differ for these countries relative to the West. Second, using Ruhm and Teague (1997) and Nickell and Layard (1999), we construct for a subset of our data, weeks of available parental leave and its square.[19] Parental leave information is available from these sources for 1981, 1991 and 1995. Since our data span 1985-94, we created an interpolated parental leave variable that incorporates some time variation. Third, again for a subset of observations for which data were available, we include a measure of the degree of occupational segregation by gender, using Anker's (1998) indexes which were based on 75 occupational titles. His indexes were collected

[18] The control for year also provides an adjustment to the price variability terms which are computed using a given year's US male sample. Year dummies in effect control for differences over time in the composition of the US male sample of workers in the ISSP.

[19] The quadratic term is included in light of our earlier discussion of the impact of short vs. long parental leave entitlements.

at one time, typically 1990 (see the data Appendix for further details). While this is not a time-varying measure, it includes a considerable amount of occupational detail, unlike published sources or the ISSP. Several of these variables may be affected by the male-female pay gap; yet each may also control for exogenous factors that influence the gender pay gap. Thus, one can make arguments for including or for excluding these additional variables. As we show below, our conclusions with respect to the effect of wage compression and net supply on the gender pay gap are robust with respect to the inclusion of these variables.[20]

In formulating our estimation strategy, we begin by noting that equation (10) combines two kinds of experiments: one is based upon between country differences in levels of the dependent and explanatory variables, and the other is based upon within country changes in these variables over time. Yet it is possible that these two sources of variation arise from different sources and hence have different consequences. Our review of the institutional literature on wage-setting suggests that there are long-standing international differences in wage-setting mechanisms that lead to varying levels of wage compression and thus gender pay gaps. While institutions may have changed in the 1980s and 1990s in a number of countries (Boeri 2000; Katz 1993), it is likely that, in most cases, the dominant factor driving changes in labor market prices within countries over time has been an increase in the demand for skilled workers due to such factors as skill-biased technological change and international trade (e.g., Freeman and Katz 1995). These technological changes included within industry demand shifts that favored white collar workers in general and thus, given traditional employment patterns, women relative to men (Berman, Bound and Griliches 1994; Katz and Murphy 1992; Blau and Kahn 1997; Welch 2000; Fortin and Lemieux 2000).

If international differences in labor market prices are primarily driven by institutional differences while within country changes are caused mostly by technological factors, time series

[20] In some cases, these variables were not available for the full set of country-year observations, in which case we omitted the observation from these particular analyses. The basic model of equation (10), with each of the three dependent variables, showed very similar results on these subsamples.

correlations of wage inequality and the gender pay gap within a country may in some sense not be comparable to cross-sectional associations between country averages of these variables. This creates an ambiguity in the empirically observed relationship between male inequality and the gender gap in the time series because our industry-based demand index will not be able to detect within-industry technological changes that have increased the demand for women. Moreover, while data limitations prevented us from formulating our demand indexes in terms of occupation-industry cells, it is unlikely that the problem could be surmounted even were the international data available at this level of disaggregation. Results of intertemporal studies within the U.S. are consistent with substantial increases in the demand for women workers within broad industry-occupation cells (Katz and Murphy 1992; Blau and Kahn 1997). An added concern regarding the time series is that changes in the unmeasured characteristics of women workers may have played a larger role in producing changes in the gender pay gap over time than in causing differences across countries. Since male inequality has tended to increase over time and the gender gap has tended to decrease, this would be another source of a possibly spurious negative correlation between male prices and the gender pay gap in the time series.

Another econometric issue in estimating equation (10) is the possibility of reverse causation. For example, as discussed above, Fortin and Lemieux (1998) posit a negative reverse causality from the gender pay gap to the level of male wage inequality. Similarly, our earlier discussion suggests that the gender pay gap is likely to negatively affect our measure of female net supply through movements along demand and supply schedules. To the extent that such reverse causation exists, OLS coefficients will be biased.

Finally, due to our use of cross-section, time-series data and the construction of our dependent and explanatory variables, there are several reasons to suspect that the classical assumptions regarding errors in least squares estimation are violated here. First, since there are multiple observations on the same country, the error terms may be correlated across years within a country. Second, the dependent variable is based on regression coefficients with their own sampling variances. Thus, the errors are likely to be heteroskedastic. Third, some of the explanatory variables are functions of regression coefficients and thus have sampling errors of their own.[21]

Because of the distinction between the sources of time series and cross-sectional variation in labor market prices as well as the econometric problems detailed above, we estimate several alternative models based on equation (10). First, as a baseline, we estimate (10) taking account of heteroskedasticity using the Huber-White procedure and allowing the error terms within a country to be correlated. Second, we estimate (10) adding country dummy variables. This is a within-country experiment in which we examine the impact of changes over time within a country. In principle, the within-country model can account for country-specific effects on the gender gap that are correlated with the basic explanatory variables. While traditional panel data models consider both between and within unit analyses as providing different estimates of the same basic parameters, our discussion above suggests that in this case they may each be measuring the effects of different phenomena.

A final econometric variation on the basic equation is to estimate reduced form models where male wage structure and female net supply are replaced by underlying variables that influence them. This portion of the analysis accounts for any reverse causality from the gender pay gap to male wage inequality. In addition to controls for occupational segregation and parental leave, we include a variety of labor market institutions, including collective bargaining coverage, minimum wages as a fraction of the average wage, unemployment insurance (UI) replacement rates, the maximum duration of UI benefit receipt, and indexes of mandated job protection for both regular and temporary employment. The specific definitions and sources of these variables are discussed in more detail below. Each of these institutional variables is likely to influence the male wage distribution by affecting the floor under union wage-setting as well as the strength of insiders' wage demands.[22] The use of collective bargaining coverage rather than

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[21] Murphy and Topel (1985) suggest a procedure for correcting standard errors in such models. We used their procedure in several of our basic models with no change in the results.

[22] For further discussion on the impact of unemployment insurance and job protection on wage-setting see Blau and Kahn (1999).

union density is important because in many countries there are pronounced differences between the two and of course it is the former that determines the labor market impact of unions. As was the case earlier in our discussion of parental leave and occupational segregation, in some cases these institutional variables are not available for our full data set.

While we employ UI system characteristics and the degree of mandated job protection in reduced form models (in addition to collective bargaining coverage), it is possible that these variables differentially affect male and female wages and thus directly affect the gender pay gap. On the one hand, men tend to be employed in more layoff prone industries (Blau and Kahn 1981). Thus, to the extent that more generous UI benefits and job protection lower the compensating differential employers in cyclically sensitive industries must pay (Topel 1984), male wage inequality and the gender pay gap may both be reduced. On the other hand, UI and job protection strengthen the power of "insiders." Given women's greater likelihood of being short-term workers, this may increase male wages more than female wages. Moreover, we make a distinction between average replacement rates and the maximum replacement rate, which is typically higher for low wage workers. The wedge between these two replacement rates is an indicator of the degree to which the UI system affects the wage structure. Finally, regulations on the use of temporary workers may be particularly relevant in affecting female labor supply.

V. Results

The means of our main dependent and explanatory variables are shown in Table 2 which gives the unweighted averages across years for each country in our sample. The table shows a large range across countries in the raw gender pay gap, which averaged from .14 in Slovenia to .85 in Japan. Apart from these extremes, the countries with the largest gaps are Switzerland, the US, Britain and Russia, while countries with relatively small gender gaps include the Eastern European countries of East Germany and Bulgaria, as well as Sweden, Italy, Ireland and New

Zealand.[23] Table 2 also shows GAP US CHARS, the gender pay gap evaluated at US values of male and female human capital characteristics (column 2), and the unexplained gender pay gap (column 3). The estimated corrected gender gaps are highly correlated across the two methods and each is also highly correlated with the total gender pay gap, likely due in part to our inability to control for actual labor market experience which has been found to be a major difference in qualifications between men and women.[24]

Table 2 also summarizes international differences in male labor market prices. As expected, the US is at or near the top in each of the inequality measures. And, while there is a positive correlation between the price effects of measured X's (STD DEV (X_{us}) or 50-10 WAGE GAP (X_{us})) and residual inequality (RESID STD DEV or 50-10 RESID GAP), a look at the rankings shows that they do appear to measure different things. For example, Australia has the 10th largest measured price variation (STD DEV) out of the 22 countries, but the 4th largest residual standard deviation; and Russia's comparison is particularly dramatic—it has the second smallest measured price variation (STD DEV) but the largest residual standard deviation.[25]

Finally the table shows net female supply relative to the US in 1989. Ireland, Italy and the Netherlands have much smaller female net supply than the other countries, a pattern driven primarily by female labor supply rather than by the demand for women. On the other hand, the Eastern European countries have the highest net supply of women, again driven primarily by female labor supply.

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[23] It is possible to keep East and West Germany separate even after unification in the ISSP data. We follow this procedure due to sizable differences between the two in the dependent and many of the explanatory variables. The country rankings here tend to be similar to those based on published data or other studies, when available (Blau and Kahn 1995; OECD various issues; ILO various issues). The gap for Australia is larger than that reported in OECD publications. This discrepancy appears to be due to the OECD data restriction to nonsupervisory employees; the gender gap we obtain from the ISSP data is consistent with studies which use other microdata for Australia (Blau and Kahn 1995; Gregory and Daly 1991). The extremely small gender gap we estimate for Italy may be in part a result of its large unreported sector (Erickson and Ichino 1995). Since the wages of many secondary workers are "off the books," respondents may be reluctant to report this income.

[24] In the regression sample of 100 country-year data points, the correlation between TOTAL GAP and GAP US CHARS is .8680; between TOTAL GAP and the UNEXPLAINED GAP, .9333; and between GAP US CHARS and the UNEXPLAINED GAP, .9085.

[25] Overall, our two measures of labor market prices are modestly positively correlated. Across the regression sample, STD DEV (X_{us}) and RESID STD DEV have a correlation coefficient of .2376, while 50-10 WAGE GAP (X_{us}) and 50-10 RESID GAP have a correlation coefficient of .3170.

Before proceeding to the results from estimating equation (10), we present figures showing the univariate relationship between one of our main dependent variable (GAP US CHARS), and two measures of wage compression: i) the 50-10 gap in predicted US male log wages using the country's own male wage equation for each year (50-10 WAGE GAP (X_{US})), shown in Figure 1; and ii) the 50-10 gap in the country's male log wage residuals for that year (50-10 RESID GAP), shown in Figure 2. All country-year data points are included and each figure includes an ordinary least squares regression line for ease of interpretation. In each case, we observe the predicted positive relationship between the gender log wage gap and male inequality. While the figures provide some suggestive evidence of an impact of institutions on the gender pay gap, we now turn to an econometric test of this relationship.

Table 3 shows regression results for the basic time-series, cross-section estimation of equation (10). Across a variety of specifications with three alternative dependent variables and two alternative measures of labor market prices, the results consistently show that higher measured prices are associated with a higher gender pay gap.[26] The coefficients for the price variables are between 1.69 and 1.85 times their standard errors in the 50-10 WAGE GAP specification and statistically significant at better than the 5% level (on a two-tailed test) in the STD DEV specification for all dependent variables. Residual inequality is also found to be positively related to the gender gap and the estimated coefficients are always larger than their standard errors; the coefficients on these variables are statistically significant at better than the 5% level (on a two-tailed test) in the GAP US CHARS specification. Note that these findings for residual inequality strongly suggest that the male regression residual contains information regarding international differences in the prices of unmeasured skills, although it may also be influenced by international differences in the quantities of unmeasured skills as well as measurement error. Finally, in each specification, higher female net supply has the expected positive effect on the gender pay gap; its coefficient is always larger than its standard error and is

[26] The results were very similar when we omitted the year dummy variables.

statistically significant at better than the 5% level (on a two-tailed test) in the GAP US CHARS and UNEXPLAINED GAP specifications. We also note that, as would be expected, the estimated effect of variation in measured prices on the gender gap tends to be larger in the GAP US CHARS specification than in the UEXPLAINED GAP specification, although these differences are never significant.

Although the three measures of the gender gap are highly correlated, the most consistently statistically significant results in Table 3 are for the specification with GAP US CHARS as the dependent variable. In this model, five of the six coefficients are significant at better than the 5% level, and the sixth is significant at the 8.7% level, all on two-tailed tests. It may be recalled that GAP US CHARS is our preferred measure since it nets out international differences in women's relative qualifications, but preserves gender differences in measured characteristics (at the US levels). The specifications in which male inequality is measured by the standard deviation tend to perform somewhat better than those using the 50-10 specification. This is perhaps not surprising in that the standard deviation takes into account inequality at all portions of the distribution and thus potentially conveys more information. The fact that the 50-10 specification performs nearly as well, despite its focus on only the bottom of the distribution, supports our expectation that it is this portion of the distribution that is particularly relevant for women.

The effects implied by the estimated coefficients in Table 3 are important in magnitude as well as generally statistically significant. To see this, it is useful to simulate the impact on the gender pay gap of going from the 25th to the 75th percentile among country averages ranked according to the variability of male measured prices, male residual inequality or female net supply. Using the GAP US CHARS dependent variable and the standard deviation inequality specification, we find that such an increase in male measured price inequality would lead to a .055 log point increase in the gender log wage gap, which is 38% of the actual 75-25 difference in the country average GAP US CHARS of .146 for our sample. A similar increase in male residual inequality is associated with an increase in the gender pay gap of .032 log points, or

22% of the actual 75-25 differential in GAP US CHARS. Finally, a 75-25 increase in female net supply would lead to a .051 log point increase in the gender pay gap, or 35% of the 75-25 differential in the pay gap. The other specifications in Table 3 usually imply effects similar in magnitude to these. These findings suggest a strong role for both labor market prices and female net supply in affecting the gender pay gap, with a quantitatively larger effect for measured prices and residuals together than for net supply.

A further indication of the quantitative importance of these variables is provided by the results presented in Table 4. The table shows, for selected Western countries, the percentage of the difference between the Western country average of GAP US CHARS and the individual country's GAP US CHARS which is accounted for by each explanatory variable.[27] Looking first at the findings for the US, we see that the variables included in the regression are more than sufficient to account for the higher gender pay gap in the US compared to the other Western countries. This is primarily due to the impact of the wage inequality variables although the higher net supply of women in the US also plays a role. The fact that the actual difference between the US and the other Western countries is less than would be expected based on the levels of the wage inequality and net supply variables implies that there are unmeasured factors favoring US women compared to women elsewhere. These may relate to women's unmeasured characteristics relative to men's and/or to laws or policies benefiting women in the US compared to their counterparts abroad. This conclusion strongly matches the results obtained in our previous work (Blau and Kahn 1992; 1995; 1996b), despite the very different methodology employed here.

The table also shows results for Sweden and Japan, two countries that are notable for their position at the extremes of the international distribution of gender gaps. The relatively low gender gap in Sweden is well explained by the male wage inequality and net supply variables.

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[27] We focus on the Western country average as the comparison under the assumption that it would be comprised of the most comparable countries. However, the results are similar when we consider the full range of countries included in the regression analysis.

Specifically, the extremely low level of male wage inequality in Sweden would be more than sufficient to account for the lower gender pay gap there, however its effect is offset somewhat by the relatively high net supply of women workers. Overall, the variables included in the regression are estimated to explain 63-67 percent of the lower gender gap in Sweden than in other Western countries. That the regression variables can explain less than the full amount of the lower gender gap in Sweden suggests that, as in the case of the US, there are unmeasured factors benefiting Swedish women relative to women elsewhere. Finally, the results for Japan indicate that the variables included in the regression analysis can explain only a relatively small proportion of the extremely large gender pay gap there: 11-12 percent. This suggests that the large gender pay gap in Japan to a considerable extent reflects unmeasured factors that are highly unfavorable to women.

In subsequent analyses we focus on specifications that employ GAP US CHARS as the dependent variable. However, before turning to these results, we consider whether the results in Table 3 suggest there are incentive effects of wage compression that affect women differently than men. To examine this question we compare the magnitude of the effect of male inequality when we control for human capital (i.e., in the GAP US CHARS and UNEXPLAINED GAP specifications) to the effect when we do not (i.e., when TOTAL GAP is the dependent variable). A larger positive effect in the former case, for example, would suggest that higher prices of human capital induce a larger investment response for women than for men—although, on net, lower inequality is associated with a smaller gender pay gap. The results are not particularly consistent with such incentive effects. No consistent pattern emerges when we look at the impact of measured prices, and the differences between the coefficients on the price variables controlling and not controlling for human capital are not statistically significant. While the effects of male residual inequality are consistently larger when we control for human capital, the differences are generally not statistically significant. Thus, the possibility that there are negative incentive effects of wage compression which are stronger for women than for men receives little empirical support here.

We found the results in Table 3 to be robust to the inclusion of the additional controls discussed above—parental leave entitlement duration and its square, a control for Eastern European countries (EAST), and the degree of occupational segregation. Further, these additional controls did not have significant effects on the gender pay gap, and interactions between EAST and the male inequality and net supply variables also proved to be insignificant individually and as a group. Moreover, the results were similar when we estimated the gender pay gap controlling for one digit occupation.

Table A1 shows that our basic findings are not due to the extreme case of the United States with its high level of wage inequality and relatively large gender pay gap.[28] The results for GAP US CHARS when the US is omitted are very similar to those for the full sample in Table 3: male inequality and female net supply both continue to have large, usually statistically significant positive effects on the gender pay gap. And the magnitudes of these effects are comparable to those reported above. Thus, while we found in earlier work (Blau and Kahn, 1992, 1995 and 1996b) that wage structure was important in explaining the differences between the gender pay gap in the US and other countries, the results presented here strongly suggest that this conclusion holds more generally.

Table A2 shows results in which we disaggregate the net supply variable into its component supply and demand indexes. The results indicate that the positive effects of net supply on the gender pay gap reported in Tables 3 are driven by female supply. In particular, female supply has positive effects on the gender pay gap that are large and statistically significant at better than the 5% level (in two tailed tests) in five of six cases. In contrast, while the effects of female demand on the gender pay gap are always negative, as expected, they are smaller than their standard errors in absolute value in each case. However, in each specification, we accept the hypothesis that the supply and demand coefficients are equal in magnitude and

[28] With the US omitted, the average values of the male inequality and gender gap variables over the remaining 21 countries are: TOTAL GAP (.303); GAP US CHARS (.265); UNEXPLAINED GAP (.278); STD DEV (.264); RESID STD DEV (.445); 50-10 WAGE GAP (.367); and 50-10 RESID GAP (.508). These may be compared to the figures for the US in Table 2.

opposite in sign (and thus that the net supply variable is an appropriate specification). The poorer showing for the demand variable may reflect our inability to control for within-industry demand.

Table 5 shows that the basic results hold up for specific subsamples: namely, wage and salary workers and married workers.[29] In all cases, the male inequality variables have positive effects on the gender pay gap, and these are usually statistically significant. Net supply continues to have a positive coefficient, although it is significant only once (the 50-10 specification for the All Workers, Married Only sample). In other specifications not shown, we obtain similar results when we include all workers but control for marital status in estimating the gender pay gap.[30]

Other alternative specifications address a potential difficulty concerning the interpretation of our price variables. Specifically, because of omitted productivity characteristics, one country may have a seemingly higher return to education than another country, a difference that would be reflected in our measured price variability term. While we would like to interpret this variable as a measure of prices, it may instead reflect the variability of unmeasured skills that are correlated with education. For example, it has been pointed out that low skill workers in the US have lower relative levels of literacy than low skill workers in other OECD countries (OECD 1998).

To take account of this possibility, we used data from the International Adult Literacy Survey 1994-6, as tabulated in OECD (1998), to form a rough control for average national levels of these unmeasured skill differences. In particular, for a subset of 11 of our 22 countries for which data were available, we computed the average difference in literacy test scores between those with upper secondary education only (i.e. middle levels) and those with less than upper secondary levels of education (lower levels).[31] This is an estimate of average differences in

[29] The former includes all wage and salary workers, regardless of marital status, while the latter includes all married workers, regardless of their self employment/wage and salary worker status.

[30] Recall that the measures of male inequality in all specifications include controls for marital status even when the measure of the gender pay gap does not include this control.

[31] The countries were Australia, Canada, Germany, Ireland, Netherlands, New Zealand, Poland, Sweden, Switzerland, the UK and the US. Since Switzerland's scores were broken down into French and German regions, we

unmeasured skills between workers with middle and low levels of education for each country. When we added this literacy variable to our basic models, the effects of prices and net supply were unaffected, and they were qualitatively similar to those for the full sample of 22 countries without the literacy variable included. Further, the literacy variable did not affect the gender pay gap. This alternative specification increases our confidence that our measured price variables really do reflect prices and that compressed returns to skills in general lead to a smaller gender pay gap.

We now present results that attempt to determine the degree to which our findings are due to differences across countries vs. changes over time within countries. Table 6 presents our basic model with country dummies added. In this case, the variation in the dependent and explanatory variables is entirely due to within country variation, while the results without country dummies (Table 3) combine the effects of within-country and between-country variation. In both the 50-10 and standard deviation specifications, the estimated effect of male measured prices on the gender pay gap remains positive but is no longer significant. Male residual inequality is now estimated to have a negative effect on the gender pay gap, but the estimated coefficients are considerably smaller than their standard errors. The effect of net supply remains positive but insignificant. Thus, the within estimates obtained by controlling for country dummies are considerably weaker and, in the case of male residual inequality, conflict with those not controlling for country dummies. This implies that the basic time-series cross-section results are primarily due to differences between countries rather than to changes over time within countries.[32]

The divergence of the fixed effects results from the estimates without controlling for country effects likely reflects the different sources of variation in male inequality in each of the two cases discussed above. An additional factor working against obtaining equally strong results

averaged them, although they were very similar to each other. The literacy variable is available only at one point in time, 1994-6.

[32] When we used country averages, we obtain similar results to those in Table 3, further underscoring the conclusion that our time series, cross section results are driven primarily by between country effects.

in the time series as in the cross-section is that most of the variation in the explanatory variables is between rather than within countries (Hamermesh, 2000).[33] Indeed, this is perhaps not surprising in that one of the attractions of international comparisons is the considerable variation in both the dependent and explanatory variables that such data provide, far exceeding what would be available for any individual country.

In light of the within country results in Table 6, the time series, cross-section results can be interpreted as primarily reflecting long-run differences across countries. One interpretation of the findings for male prices is that wage-setting institutions such as collective bargaining contracts that call for high wage floors lower male inequality and the gender pay gap as well. Further, where women are scarce in relation to the demand for labor in industries that traditionally employ them, the gender pay gap will be lower. This finding could reflect basic supply and demand in the labor market or selectivity of the female labor force. But these interpretations can only be made subject to the usual caveat concerning cross-sectional differences across countries: there may be factors we cannot control for that influence male labor market prices and female net supply as well as the gender pay gap. The reduced form results presented in Tables 7-9 shed additional light on this issue and also respond to the concern that wage inequality and possibly female net supply are potentially endogenous.

Specifically, in Tables 7-9, we present a number of reduced form specifications for the determination of the gender pay gap, omitting our measures of male wage inequality and instead including information on institutions that are expected to affect wage inequality in general or the gender gap specifically. These institutions include collective bargaining coverage, minimum wage laws, unemployment insurance systems, and job protection, as well as parental leave entitlements. The Data Appendix contains details on how these variables were constructed. As much as possible, we have constructed time varying measures of these institutions, to reflect the ongoing reforms of labor market institutions in the OECD (Boeri 2000). This analysis was based

[33] Seventy-five percent or more of the variation in the male inequality variables and 95 percent of the variation in net supply is associated with between country variation.

in most cases on the 14 countries for which we had this information.[34] There has been some relaxation of employment protection restrictions and some reduction in collective bargaining coverage in several countries over the 1980s and 1990s, and this is reflected in our institutional measures. However, as was the case with our wage compression and net supply measures, the cross country variation in these institutions is much greater than their variation over time within countries. Specifically, between 89 and 99.96 percent of the total variation in collective bargaining coverage, unemployment insurance rules, parental leave entitlements, and job protection in our sample occurs between countries. Therefore, while there is time variation in these measures of labor market institutions, we are essentially looking at the effects of cross-sectional differences in them.

Table 7 shows results for including collective bargaining coverage alone (with year dummies) and with our original measures of wage compression and female net supply. With no covariates besides year dummies, collective bargaining coverage has a coefficient of -.0029 that is highly significant. This coefficient corresponds to a large impact of collective bargaining on the gender pay gap, implying that an increase in collective bargaining coverage from the 25th to the 75th percentile in our sample of 14 countries (i.e., going from 47 to 82 percent coverage) decreases the gender pay gap by .10 log points. This predicted fall is larger than the actual 75-25 difference in the gender pay gap among these countries of .07 log points. The second and third specifications in Table 7 reproduce the corresponding ones in Table 3 for the sample of countries for which information on collective bargaining is available. As may be seen in the table, we obtain similar results to those in Table 3 for this subsample: measured male price inequality, male residual wage inequality, and female net supply all have positive, significant effects on the gender pay gap.

The final two specifications in Table 7 show that there is still a direct effect of collective bargaining even when wage compression and female net supply are included in the regression.

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[34] The countries were Austria, Australia, Britain, Canada, West Germany, Ireland, Italy, Japan, the Netherlands, New Zealand, Norway, Sweden, Switzerland, and the US.

In these instances, the collective bargaining coefficient is -.0024 to -.0028 and is 2-2.2 times its standard error in absolute value. A possible reason for this finding is that countries with very high levels of collective bargaining coverage have a greater coverage of female-dominated sectors of the labor market than countries with lower overall levels of collective bargaining. A further possibility is that the estimated direct effect of collective bargaining may be due in part to the fact that the extent of wage inequality is not perfectly captured by our inequality variables. Our estimate of the impact of measured prices in based on just a few variables reflecting education, potential experience and marital status, and our measure of residual inequality reflects not only unmeasured prices but also differences across countries in the variation in unmeasured quantities and measurement error. Thus the extent of collective bargaining may continue to reflect the impact of wage compression, even when our inequality variables are included. When we control for collective bargaining, measured inequality in male prices and female net supply are still positively associated with the gender pay gap with effects that are 1.6-2.0 (measured price inequality) or 1.4-1.9 (female net supply) times their standard errors. However, in contrast to the results in Table 3, there is no longer a positive association between male residual wage inequality and the gender pay gap; this effect is negative and insignificant in Table 7.

Table 8 provides further evidence on the importance of institutions in affecting the gender pay gap by looking at both minimum wages and collective bargaining. The first column shows the impact of the minimum wage level (relative to the average wage) when only year dummies included. The coefficient is significantly negative, suggesting a negative correlation between the "bite" of minimum wages and the gender pay gap.[35] However, when we control for collective bargaining coverage and minimum wages at the same time, the effect of minimum wages becomes much smaller and is now insignificant. The effect of collective bargaining, however, remains significantly negative and large in absolute value even controlling for minimum wages. We further note that in several countries where there is a national wage floor,

[35] The sample in Table 8 includes one less observation than that in Table 7 because we lack data on minimum wages in Switzerland.

this is set in collective bargaining agreements covering most of the labor force (in our sample, this includes Austria, Norway and Sweden) or in the application of collective bargaining agreements to nonunion workers (for example, Germany, Italy and Australia). In other countries, however, the government actually passes minimum wage legislation (e.g., Netherlands, UK, Japan, New Zealand, Canada, US, and Ireland).[36] However, even where there is a statutory minimum, its level may be influenced by the strength of unions in influencing the political process. It may therefore be difficult to disentangle the impact of legislated minima from the effect of collective bargaining.

Bearing these difficulties in mind, the last two specifications shown in Table 8 attempt to more finely distinguish between the effects of collective bargaining and minimum wage legislation on the gender pay gap. In the first of these, we add a dummy variable for the existence of a statutory minimum. While this dummy is not significant, and the minimum wage variable remains insignificant, in both cases the effect is negative and 1.4-1.5 times its standard error in absolute value. Note that the impact of collective bargaining coverage remains highly significant and large in absolute value. The last model adds an interaction term between having a statutory minimum and the minimum wage (relative to the average). In this case, the main effect for the minimum/average becomes statistically significant, although an F-test shows that the effect of the minimum/average for countries with statutory minima (i.e., the sum of the minimum/average main effect and its interaction with the statutory minimum dummy) is not significant. The impact of collective bargaining, as always, remains significantly negative and large in absolute value. This model suggests that a national wage floor has a significantly negative effect on the gender pay gap if it comes about through collective bargaining, but not if it comes about through legislation. However, this result may be misleading. During our sample period, in both the UK and Ireland, the statutory minima applied only to certain industries and covered a very small portion of the labor force, suggesting that, in effect, these countries did not

[36] See Nickell and Layard (1999) and Neumark and Wascher (1999) for further discussion of minimum wage systems.

have a national minimum wage (Nickell and Layard 1999). When we re-estimated the model giving these countries a zero for the minimum/average and the statutory minimum dummy variables, the minimum/average variable now had insignificant effects on the gender pay gap for all countries. The impact of collective bargaining coverage, however, remained negative, significant and large in absolute value. Overall, then, it appears that collective bargaining coverage has a stronger effect on the gender pay gap than minimum wage laws do. This conclusion makes sense if collective bargaining has larger wage effects at the bottom than statutory minimum wage laws.[37]

Table 9 shows results for collective bargaining, controlling for the effect of parental leave, unemployment insurance and job protection policies and, in one specification, occupational segregation and female net supply as well. In all cases, collective bargaining continues to have a negative, significant effect on the gender pay gap that is large in absolute value. Evidence on the effects of the other policies was sensitive to the inclusion of the occupational segregation and female net supply variables. When these variables are not included (i.e., in the first and second specifications shown in the table), the labor market policies variables have insignificant effects both individually and as a group. However, controlling for net supply and segregation, these policy variables are significant at the 8% level as a group, although not individually.[38] Finally, the segregation index has the expected positive effect on the gender pay gap, although it is only 1.53 times its standard error.

These reduced form results for collective bargaining further strengthen our confidence in our basic findings for male wage inequality. Nonetheless it is always possible that there are omitted variables that are correlated with collective bargaining coverage and the gender pay

[37] Additionally, there was no evidence that the effect of collective bargaining coverage was significantly affected by the relative level of minimum wages.

[38] This difference in the impact of the policy variables depending on whether we control for net supply and segregation is not due to the smaller sample (66 observations) for which segregation data are available. When we re-estimated the second specification shown in Table 9 on this subsample, the policy variables remained jointly and individually insignificant.

gap.[39] For example, there are other social programs that differ in generosity across countries which could potentially affect both male inequality and the gender gap. For example, we do not have good data on the enforcement of equal employment opportunity laws, and variation in this factor may help explain differences in the gender pay gap. One possible indicator for such enforcement is the incidence of employment in very small firms, many of which are effectively exempt from labor market regulation (Carrington, McCue and Pierce 2000). We examined this issue for the nine countries for which the OECD (1994) collected information during the 1990-2 period on the fraction of employment in firms with less than 20 workers.[40] While the sample size fell to 47 in this analysis, we still found a negative, significant effect of collective bargaining on the gender pay gap, while the effect of the incidence of employment in small firms was small and insignificant. Although the incidence of employment by firm size can clearly be affected by relative wages and by labor market institutions, the impact of collective bargaining was nonetheless robust with respect to this possible measure of the degree to which equal employment opportunity laws are enforced.

The results discussed above indicate that our findings for collective bargaining coverage are strong whether or not controls for other institutions are included. Further, according to Summers, Gruber and Vergara (1993), it is centralizing wage-setting institutions that are primarily responsible for the growth of generous social programs in countries like Sweden. If this is the case, then our models can be viewed as reduced forms of such a process in which the prime mover is collective bargaining.

[39] While we have used the degree of collective bargaining coverage, it is also possible that the degree of coordination of collective bargaining can have an independent effect on wage inequality and thus on the gender pay gap. We investigated the impact of coordination using Nickell's (1997) 6 point scale, which improves upon Calmfors and Driffill's (1988) original centralization measure by taking into account pattern bargaining and coordination. Unfortunately, in the sample of 14 countries for which we had data, collective bargaining coverage and coordination were too closely correlated to reliably disentangle their separate effects; the two variables had a correlation coefficient of .67. Collective bargaining coverage and Calmfors and Driffill's (1988) original centralization measure were correlated at 0.71 with country rankings for collective bargaining coverage (OECD 1997, p. 73).

[40] The countries were Australia, Canada, Germany, Italy, Japan, Sweden, Switzerland, the UK and the United States.

Another potential omitted variable would be social norms. For example, it is possible that a high degree of collective bargaining coverage, low male inequality, and a small gender gap all reflect an underlying social egalitarianism. While this is plausible, it seems unlikely to provide a full explanation for our findings. For one thing, there is considerable evidence that discontinuous changes in wage setting institutions within countries do produce changes in wage inequality (Blau and Kahn 1999). For another, there are the examples of countries like Austria and Italy that have highly centralized wage setting and relatively low male inequality and gender differentials and yet show little evidence of favorable treatment of women as a group otherwise.

A final consideration relates to a possible alternative interpretation of our finding of a negative association between wage compression and the gender pay gap, at least some of which appears to be due to the impact of collective bargaining coverage. We believe that these results indicate that bringing up the bottom of the pay distribution raises the relative wages of low-paid workers such as women. However, it is also possible that the negative relationship between pay compression and the gender pay gap reflects a positive impact of high wage floors on women's labor market skills, particularly their levels of actual labor market experience, although it may be recalled that the sign of this relationship is theoretically ambiguous. Some indication of the importance of this effect can be gained from panel data sets available for the US, Sweden and Germany which, unlike the ISSP, do have information on actual experience.

Germany and Sweden are both highly unionized societies with 80-90% collective bargaining coverage, in contrast to the US, with its 18% coverage rate in 1990. And Table 2 shows that the male wage distribution is indeed much more compressed in these countries than in the US. Using data sets we employed in earlier work (Blau and Kahn 1996b and 2000), we find that the male-female gap in actual experience in 1984 was about 6 years in the US and 5 years in Sweden and Germany. While these figures are consistent with the idea that high wage floors encourage female labor force commitment, in fact female relative labor force participation rates among these three countries in the 1980s were highest in Sweden but higher in the US than in

West Germany.[41] And the male-female experience gap in the US declined steadily over the 1980s (Blau and Kahn 1997). Further, US women are much more likely to work full-time than women in these other countries (Blau and Kahn 1995). Thus, by the end of the 1980s, it is possible that US women had higher relative experience levels than women in several more highly unionized countries. It is therefore unlikely that a union effect on women's unmeasured labor market qualifications is responsible for our basic findings. Even if it were, however, one could still attribute an ultimate causal role to institutions of wage compression.

VI. Conclusions

This paper tested the hypotheses that overall wage compression and low female supply relative to demand reduce a country’s gender pay gap. Using micro-data for 22 countries over the 1985-94 period, we found that more compressed male wage structures and lower female net supply are both associated with a lower gender pay gap. Since it is likely that labor market institutions are responsible for an important portion of international differences in wage inequality, the inverse relationship between the gender pay gap and male wage inequality suggests that wage-setting mechanisms, such as encompassing collective bargaining agreements, that provide for relatively high wage floors raise the relative pay of women, who tend to be at the bottom of the wage distribution. Consistent with this view, we find that the extent of collective bargaining coverage in each country is significantly negatively related to the gender gap. Overall, our results provide strong evidence that wage setting institutions have important effects on the gender pay gap, and some evidence of the impact of the market forces of supply and demand as well.

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[41] Specifically, over the 1985-88 period, the ratio of the female to the male labor force participation rate was 58.6% in West Germany, 73.3% in the US, and 94.5% in Sweden (Blau and Ferber 1992). The experience gap accounted for only about .03 log points of the 1984 .21 log point difference in the Swedish-US gender gaps (Blau and Kahn 1996b).

Data Appendix

In this appendix, we briefly describe the sources for the construction of the institutional variables used in Tables 7-9:

Collective bargaining coverage was taken from OECD (1997), which had data for 1980, 1990 and 1994. We used linear interpolation to create a time series for this variable. For several countries, 1980 information was not available. In these cases, we used the 1990 figures. For Ireland, we assigned a value of 75 for collective bargaining coverage, in light of Layard, Nickell and Jackman's (1991) classification of this country as having coverage in the "70 percent and above" range.

Parental leave was constructed from Ruhm and Teague (1997), who reported 1981 and 1991 values for parental leave and Nickell and Layard (1999), who reported 1995 values for this variable. Again, we used linear interpolation to create a time series.

Regular and temporary employment protection were taken from OECD (1999), which reported indexes for protection for the late 1980s and the late 1990s. We used the late 1980s values for the years 1985-1989, and the late 1990s values for the years 1990-94. Results were similar when we used a simple average of the late 1980s and the late 1990s for the 1990-94 period. The scales for these variables ranged from 0 to 4, with a higher value signifying more government restrictions on firms' use of labor.

Unemployment insurance (UI) benefit duration was taken from Nickell and Layard (1999) and is a one-time average for 1989-94. The average and maximum UI replacement rates were taken from Blanchard and Wolfers (2000). These authors compute 1985-89 and 1990-94 averages for these variables. We use the former for 1985-89 and the latter for 1990-94, thus creating a time-varying variable for UI replacement rates.

Minimum wages/average wages are a one-time average taken from Nickell and Layard (1999) and refer to 1991-4.

The Segregation Index comes from Anker (1998), Table 9.1. It is based on 75 nonagricultural two-digit occupations. Data refer to: 1990 for Australia, Austria, Canada, the

Hungary, Japan, Netherlands, Norway and the UK; 1991 for Sweden and the US; 1989 for W. Germany; 1986 for New Zealand; 1980 for Switzerland; 1985 for Bulgaria; 1988 for Poland; 1981 for Italy and Slovenia (former Yugoslavia).

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Figure 1: Gender Log Wage Gap at US Male and Female X's by 50-10 Gap in Measured Male Human Capital Prices

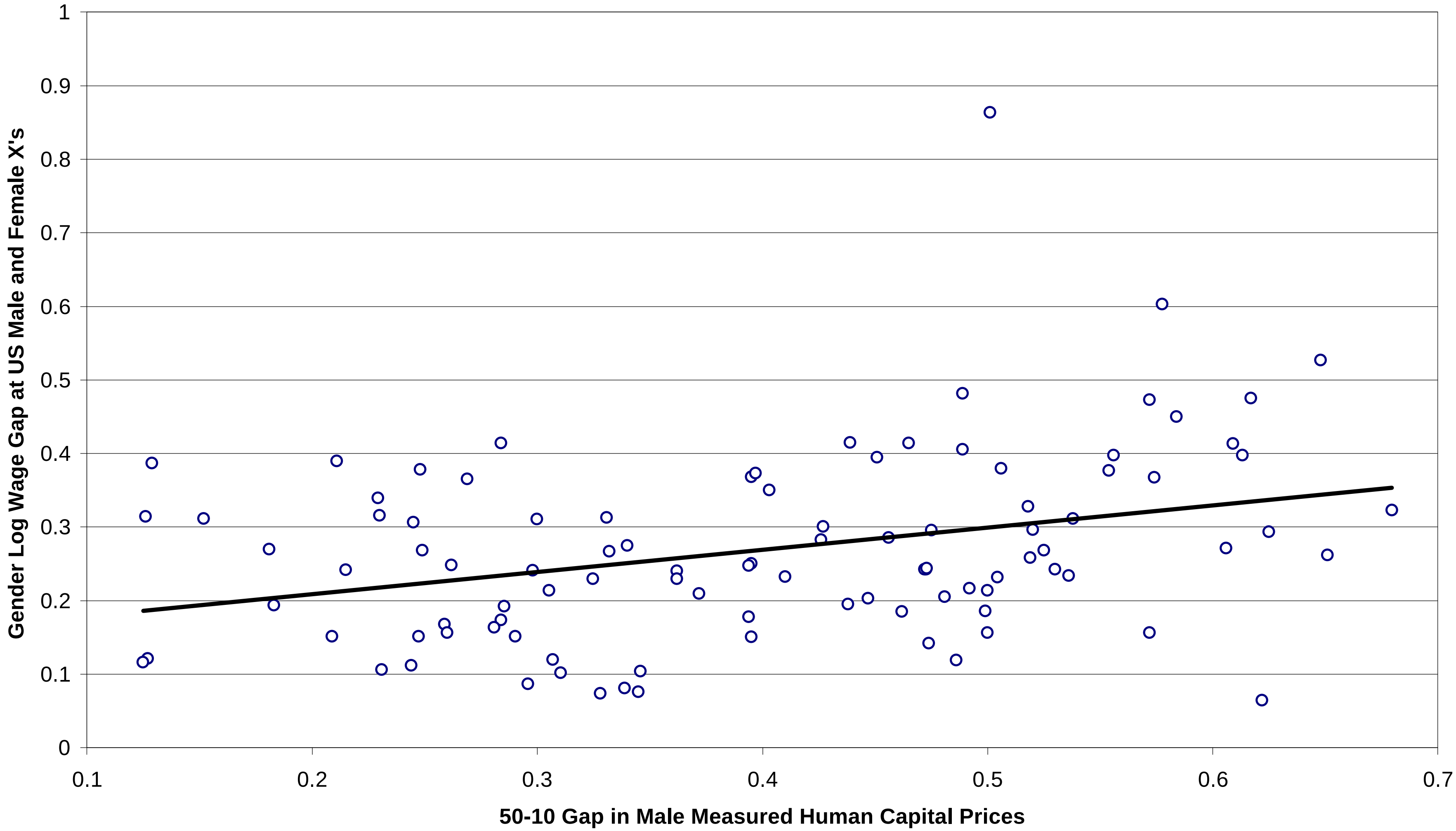

Figure 2: Gender Log Wage Gap at US Male and Female X's by 50-10 Gap in Male Log Wage Residuals

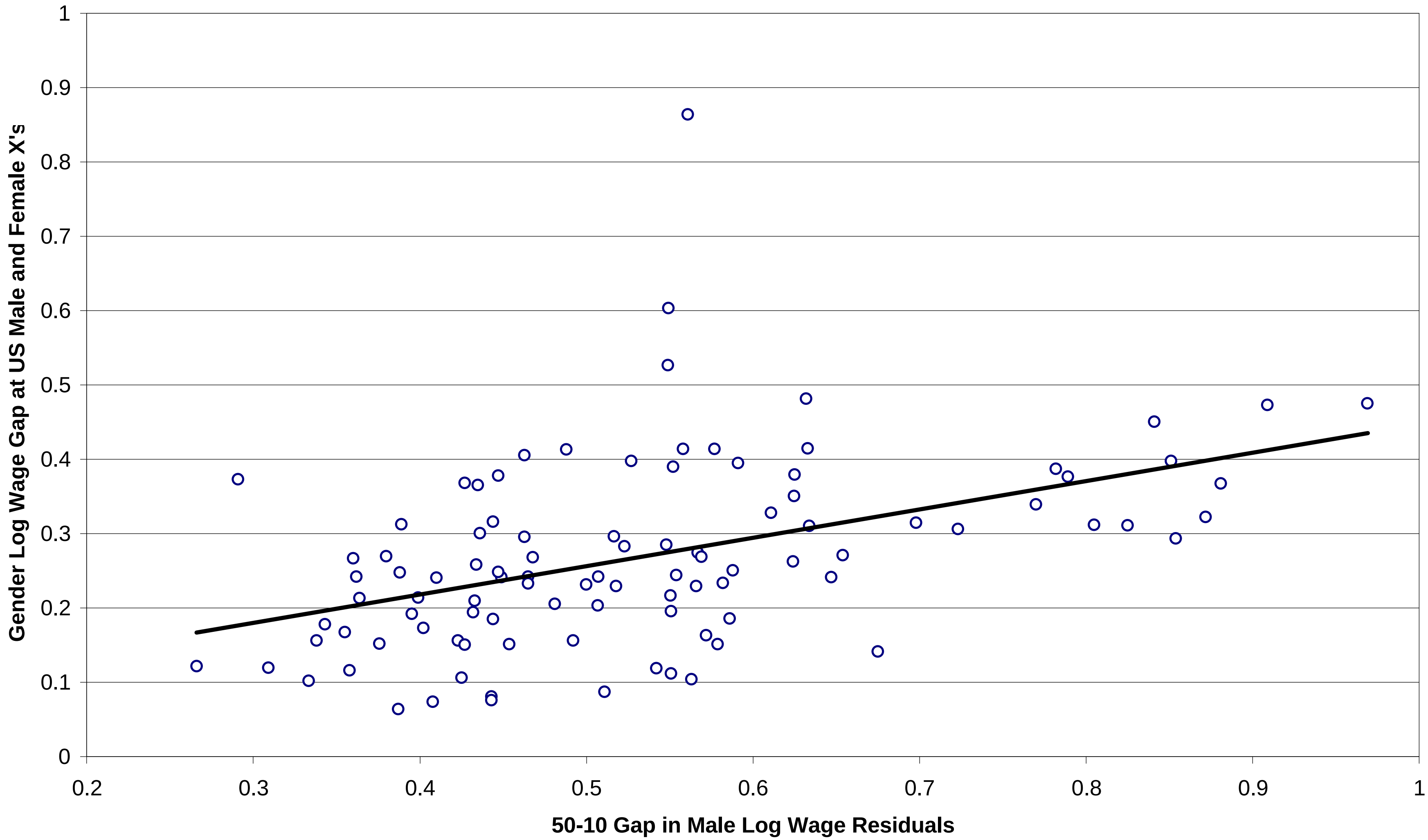

Table 1: Countries, Years, and Earnings Concepts in the ISSP Data

Country	Years of Data	Earnings Concept
Australia	86,87,90,91,94	Annual Gross Wage and Salary Income
Austria	85-87,89,91-92,94	Monthly Net Earnings
Britain	85-94	Annual Gross Earnings
Bulgaria	92-93	Monthly Earnings
Canada	92-94	Annual Gross Personal Income
Czech Rep	92,94	Monthly Net Income
East Germany	90-93	Monthly Net Earnings
West Germany	85-93	Monthly Net Earnings
Hungary	88-94	Monthly Gross Earnings
Ireland	88-90,93-94	Annual Net Earnings (88-90); Weekly Net Earnings (93-94)
Israel	93-94	Monthly Earnings
Italy	86,88,90,92-94	Monthly Net Income
Japan	93-94	Annual Gross Earnings
Netherlands	88-89	Annual Net Earnings
New Zealand	91-94	Annual Gross Income
Norway	89-94	Annual Gross Earnings
Poland	91-94	Monthly Net Earnings, Main Job
Russia	91-94	Monthly Net Earnings
Slovenia	91-94	Monthly Regular Income
Sweden	94	Monthly Gross Earnings
Switzerland	87	Monthly Net Income
USA	85-94	Annual Gross Earnings

Table 2

Means of Dependent and Explanatory Variables by Country*

	TOTAL GAP	GAP US CHARS	UNEXPLAINED GAP	STD DEV (X_{us})	RESID STD DEV	50-10 WAGE GAP (X_{us})	50-10 RESID GAP	NET SUPPLY
Australia	0.305	0.267	0.266	0.286	0.536	0.418	0.547	-0.054
Austria	0.290	0.199	0.251	0.311	0.404	0.415	0.439	0.010
Britain	0.366	0.348	0.373	0.442	0.437	0.515	0.548	-0.025
Bulgaria	0.179	0.233	0.216	0.158	0.498	0.168	0.639	0.265
Canada	0.282	0.224	0.239	0.297	0.418	0.472	0.535	-0.004
Czech Rep	0.309	0.293	0.299	0.163	0.399	0.205	0.412	0.181
E. Germany	0.178	0.121	0.173	0.141	0.323	0.213	0.356	0.123
W. Germany	0.320	0.222	0.280	0.310	0.363	0.501	0.416	-0.062
Hungary	0.291	0.263	0.275	0.224	0.399	0.269	0.449	0.147
Ireland	0.220	0.228	0.211	0.347	0.480	0.499	0.564	-0.181
Israel	0.314	0.341	0.341	0.222	0.420	0.266	0.564	-0.088
Italy	0.228	0.157	0.190	0.220	0.438	0.323	0.459	-0.191
Japan	0.852	0.733	0.789	0.386	0.462	0.539	0.555	-0.013
Netherlands	0.274	0.196	0.171	0.272	0.352	0.447	0.354	-0.203
New Zealand	0.188	0.182	0.196	0.222	0.481	0.340	0.541	0.018
Norway	0.275	0.263	0.257	0.245	0.357	0.378	0.409	0.013
Poland	0.292	0.345	0.345	0.313	0.556	0.347	0.654	0.239
Russia	0.364	0.357	0.361	0.144	0.702	0.180	0.727	0.265
Slovenia	0.144	0.125	0.150	0.228	0.499	0.259	0.470	0.199
Sweden	0.216	0.214	0.223	0.216	0.354	0.305	0.399	0.048
Switzerland	0.479	0.262	0.240	0.407	0.461	0.651	0.624	-0.110
United States	0.395	0.395	0.385	0.408	0.678	0.579	0.842	-0.004

*Unweighted averages across years for each country.

Table 3: Basic Time Series-Cross Section Regression Results

	DEPENDENT VARIABLE: TOTAL GAP				DEPENDENT VARIABLE: GAP US CHARS (Bjm*Xum - Bjf*Xuf)				DEPENDENT VARIABLE: UNEXPLAINED GAP (Bjm*Xjf - Bjf*Xjf)			
	Coeff.	Std. Err.	Coeff.	Std. Err.	Coeff.	Std. Err.	Coeff.	Std. Err.	Coeff.	Std. Err.	Coeff.	Std. Err.
50-10 PREDICTED WAGE GAP (X_{us})	0.3448	0.1862			0.3053	0.1701			0.2961	0.1752		
MALE RESID 50-10	0.1097	0.0818			0.2573	0.0949			0.1741	0.0798		
PREDICTED WAGE STD DEV (X_{us})			0.4138	0.1928			0.4850	0.1512			0.4382	0.1707
MALE RESID STD DEV			0.1049	0.0997			0.2750	0.0937			0.1389	0.0966
NET SUPPLY	0.2188	0.1339	0.1233	0.1167	0.2781	0.1252	0.2132	0.1009	0.2865	0.1327	0.2282	0.1101
Year dummies	yes		yes		yes		yes		yes		yes	
sample size	100		100		100		100		100		100	
R2	0.2641		0.2284		0.3310		0.3268		0.2358		0.2102	

Notes: Gender gap computed from equations including education, potential experience and its square. Male prices and residuals are from equations that additionally include marital status. Standard errors are heteroskedasticity-robust and are corrected for within-country error correlation. TOTAL GAP is the gender log difference in hours-corrected earnings; GAP US CHARS is the gender pay gap assuming US male and female values for the X variables; UNEXPLAINED GAP is the difference in a country's predicted pay for women based on the country's male and its female wage coefficients, evaluated at female means.

Table 4: Accounting for the Difference between Individual Country Gender Pay Gaps and the Average for the Western Nations: Selected Countries

Dependent Variable: GAP US CHARS

	Specification	
	STD DEV	50-10 GAP
US vs. Other Western Countries: $(D_j - D_w) = .1209$		
% of Difference Due to:		
Measured Prices	44.0	36.7
Residual Inequality	57.3	73.6
Net Supply	9.9	13.0
Total	111.2	123.3
Sweden vs. Other Western Countries: $(D_j - D_w) = -.0734$		
% of Difference Due to:		
Measured Prices	63.6	61.4
Residual Inequality	35.7	43.6
Net Supply	-32.4	-42.3
Total	66.9	62.7
Japan vs. Other Western Countries: $(D_j - D_w) = .4836$		
% of Difference Due to:		
Measured Prices	8.6	6.5
Residual Inequality	1.2	2.0
Net Supply	2.1	2.7
Total	11.9	11.2

Note: D_i and D_w are respectively the average gender log wage gap (GAP US CHARS) in country j and in Western countries.

Table 5: Time Series-Cross Section Regression Results for Subsamples

Dependent Variable: GAP US CHARS

	Wage and Salary Workers Only				**All Workers, Married Only**			
	Coeff.	Std. Err.	Coeff.	Std. Err.	Coeff.	Std. Err.	Coeff.	Std. Err.
50-10 PREDICTED WAGE GAP (X_{us})	0.1192	0.1791			0.6390	0.1701		
MALE RESID 50-10	0.3190	0.0953			0.2322	0.0785		
PREDICTED WAGE STD DEV (X_{us})			0.2841	0.1653			0.8060	0.1939
MALE RESID STD DEV			0.3142	0.1109			0.3066	0.1114
NET SUPPLY	0.1155	0.1192	0.1086	0.0966	0.3281	0.1319	0.1437	0.1301
Year dummies	yes		yes		yes		yes	
sample size	100		100		100		100	
R2	0.2793		0.2503		0.3993		0.3768	

Notes: Gender gap computed from equations including education, potential experience and its square. Male prices and residuals are from equations that additionally include marital status and are estimated across all workers. Standard errors are heteroskedasticity-robust and are corrected for within-country error correlation. GAP US CHARS is the gender pay gap assuming US male and female values for the X variables.

Table 6: Basic Time Series-Cross Section Regression Results With Country Effects

Dependent Variable: GAP US CHARS, Sample Includes All Workers

	Coeff.	Std. Err.	Coeff.	Std. Err.
50-10 PREDICTED WAGE GAP (X_{us})	0.1276	0.1431		
MALE RESID 50-10	-0.0651	0.1639		
PREDICTED WAGE STD DEV (X_{us})			0.1480	0.2078
MALE RESID STD DEV			-0.0391	0.1903
NET SUPPLY	0.3700	0.6908	0.3558	0.7067
Year dummies	yes		yes	
Country Dummies	yes		yes	
sample size	100		100	
R2	0.7754		0.7739	

Notes: Gender gap computed from equations including education, potential experience and its square. Male prices and residuals are from equations that additionally include marital status. Standard errors are heteroskedasticity-robust and are corrected for within-country error correlation. GAP US CHARS is the gender pay gap assuming US male and female values for the X variables.

Table 7: Results for Collective Bargaining Coverage and Wage Compression

Dependent Variable: GAP US CHARS

	Coeff.	Std. Err.	Coeff.	Std. Err.	Coeff.	Std. Err.	Coeff.	Std. Err.	Coeff.	Std. Err.
COLLECTIVE BARGAINING COVERAGE	-0.0029	0.0007					-0.0028	0.0013	-0.0024	0.0012
50-10 PREDICTED WAGE GAP (X_{us})			0.3638	0.2184			0.2780	0.1753		
MALE RESID 50-10			0.1937	0.1178			-0.1323	0.2085		
PREDICTED WAGE STD DEV (X_{us})					0.5344	0.1786			0.3519	0.1722
MALE RESID STD DEV					0.2447	0.1009			-0.0878	0.2318
NET SUPPLY			0.4274	0.1748	0.3803	0.1673	0.2283	0.1200	0.1883	0.1336
Year dummies	yes		yes		yes		yes		yes	
sample size	71		71		71		71		71	
R2	0.4546		0.3942		0.4140		0.5013		0.5063	

Notes: Gender gap computed from equations including education, potential experience and its square. Male prices and residuals are from equations that additionally include marital status and are estimated across all workers. Standard errors are heteroskedasticity-robust and are corrected for within-country error correlation. GAP US CHARS is the gender pay gap assuming US male and female values for the X variables.

Table 8: Results for Collective Bargaining Coverage and Minimum Wages

Dependent Variable: GAP US CHARS

	Coeff.	Std. Err.	Coeff.	Std. Err.	Coeff.	Std. Err.	Coeff.	Std. Err.
MINIMUM WAGE/AVERAGE WAGE	-0.6426	0.1843	-0.1674	0.1912	-0.3167	0.2171	-0.3604	0.1297
COLLECTIVE BARGAINING COVERAGE			-0.0024	0.0008	-0.0030	0.0011	-0.0032	0.0016
STATUTORY MINIMUM DUMMY					-0.0741	0.0495	-0.1482	0.3100
STATUTORY MINIMUM DUMMY * (MINIMUM WAGE/AVERAGE WAGE)							0.1345	0.5902
Year dummies	yes		yes		yes		yes	
sample size	70		70		70		70	
R2	0.3927		0.4634		0.4834		0.4840	

Notes: Gender gap computed from equations including education, potential experience and its square. Male prices and residuals are from equations that additionally include marital status and are estimated across all workers. Standard errors are heteroskedasticity-robust and are corrected for within-country error correlation. GAP US CHARS is the gender pay gap assuming US male and female values for the X variables. MINIMUM WAGE/AVERAGE WAGE refers to nationwide minimum pay level, whether it is set by law or in collective bargaining. STATUTORY MINIMUM DUMMY=1 if country has a law mandating a particular minimum wage level.

Table 9: Further Results for Collective Bargaining Coverage and Other Institutions

Dependent Variable: GAP US CHARS

	Coeff.	Std. Err.	Coeff.	Std. Err.	Coeff.	Std. Err.
COLLECTIVE BARGAINING COVERAGE	-0.0040	0.0017	-0.0041	0.0023	-0.0054	0.0026
PARENTAL LEAVE (WEEKS)	0.0017	0.0030	0.0015	0.0024	-0.0006	0.0032
PARENTAL LEAVE SQUARED(/100)	-0.0004	0.0016	-0.0004	0.0015	0.0013	0.0019
UI BENEFIT DURATION			0.0017	0.0167	0.0011	0.0279
AVE UI REPLACEMENT RATE			-0.0035	0.0037	-0.0059	0.0040
MAX UI REPLACEMENT RATE			0.0014	0.0022	0.0011	0.0026
JOB PROTECTION: REGULAR EMPLOYEES			0.0458	0.0748	0.0302	0.0800
JOB PROTECTION: TEMPORARY EMPLOYEES			-0.0137	0.0241	0.0344	0.0303
SEGREGATION INDEX					1.2553	0.8227
NET FEMALE SUPPLY					0.1915	0.4259
Year dummies	yes		yes		yes	
Sample Size	71		71		66	
R2	0.4872		0.5202		0.5779	

Notes: Gender gap computed from equations including education, potential experience and its square. Male prices and residuals are from equations that additionally include marital status and are estimated across all workers. Standard errors are heteroskedasticity-robust and are corrected for within-country error correlation. GAP US CHARS is the gender pay gap assuming US male and female values for the X variables.

Table A1: Basic Time Series-Cross Section Regression Results, United States Excluded

Dependent Variable: GAP US CHARS

	Coeff.	Std. Err.	Coeff.	Std. Err.	Coeff.	Std. Err.	Coeff.	Std. Err.
50-10 PREDICTED WAGE GAP (X_{us})	0.3391	0.1891			0.2924	0.1425		
MALE RESID 50-10	0.3311	0.1200			0.3414	0.1359		
PREDICTED WAGE STD DEV (X_{us})			0.4960	0.1717			0.5431	0.1818
MALE RESID STD DEV			0.2657	0.1336			0.0878	0.1740
NET SUPPLY	0.2733	0.1301	0.2028	0.1038	0.1369	0.1542	0.1877	0.1318
Occupational Controls	no		no		yes		yes	
Year dummies	yes		yes		yes		yes	
sample size	90		90		83		83	
R2	0.2717		0.2511		0.3840		0.3258	

Notes: In models without occupational controls, gender gap is computed from equations including education, potential experience and its square; in these models, males prices and residuals are from equations that also include marital status. In the last two columns (with occupational controls), the gender gap and male prices are from equations that additionally include one digit occupational dummies. Standard errors are heteroskedasticity-robust and are corrected for within-country error correlation.

Table A2: Time Series-Cross Section Regression Results, United States Included, Separate Female Supply and Demand Effects

	DEPENDENT VARIABLE: TOTAL GAP				DEPENDENT VARIABLE: GAP US CHARS (Bjm*Xum - Bjf*Xuf)				DEPENDENT VARIABLE: UNEXPLAINED GAP (Bjm*Xjf - Bjf*Xjf)			
	Coeff.	Std. Err.	Coeff.	Std. Err.	Coeff.	Std. Err.	Coeff.	Std. Err.	Coeff.	Std. Err.	Coeff.	Std. Err.
50-10 PREDICTED WAGE GAP (X_{us})	0.3321	0.2331			0.2731	0.2103			0.2726	0.2188		
MALE RESID 50-10	0.1092	0.0822			0.2560	0.0932			0.1731	0.0794		
PREDICTED WAGE STD DEV (X_{us})			0.3797	0.2524			0.4276	0.2084			0.3948	0.2283
MALE RESID STD DEV			0.1114	0.0991			0.2858	0.0912			0.1471	0.0920
FEMALE SUPPLY	0.2369	0.1375	0.1663	0.1462	0.3242	0.1012	0.2856	0.1009	0.3202	0.1259	0.2829	0.1248
FEMALE DEMAND	-0.1613	0.3271	-0.0084	0.2559	-0.1319	0.3102	-0.0195	0.2503	-0.1797	0.3162	-0.0818	0.2508
F-test signif. Level:												
H_0: $b_{femdem}+ b_{femsup}=0$	0.8346		0.6431		0.5344		0.3623		0.6764		0.5147	
Year dummies	yes		yes		yes		yes		yes		yes	
sample size	100		100		100		100		100		100	
R2	0.2651		0.2329		0.3372		0.3388		0.2395		0.2179	

Notes: Gender gap computed from equations including education, potential experience and its square. Male prices and residuals are from equations that additionally include marital status. Standard errors are heteroskedasticity-robust and are corrected for within-country error correlation. RAW GAP is the gender log difference in hours-corrected earnings; US GAP is the gender pay gap assuming US male and female values for the X variables; UNEXPLAINED GAP is the difference in a country's predicted pay for women based on the country's male and its female wage coefficients and the female means.