Gender Differences in the Labor Market Effects of the Dollar

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Abstract

Although the dollar has been shown to influence the expected wages of workers, the analysis to date has focused on the male workforce. We show that exchange rate fluctuations also have important implications for women's wages. The dominant wage effects for women—like those for men—arise at times of job transitions. Changes in the value of the dollar can cause the wage gap between women who change jobs and women who stay on in their jobs to expand or contract sharply, with the most pronounced effects occurring among the least educated women and women in highly competitive manufacturing industries. In addition, it appears that women who stay on in their jobs show greater wage sensitivity to currency movements than do their male counterparts.

JEL codes: F31, F3, F4, J30, E24

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by Linda Goldberg and Joseph Tracy

Dollar fluctuations are a regular part of the U.S. economic landscape. Recent research has shown that these fluctuations can have important effects on investment, stock prices, traded goods prices, and specifically for labor markets in the United States. While the aggregate net consequences for employment appear to be small, the wage implications can be sizable. These wage consequences also are concentrated in certain populations of workers — those located in more competitive industries and those with less education. Using a sample of male workers, Goldberg and Tracy (2001) show that most of the wage responsiveness to dollar fluctuations occurs for workers at times of job transitions, and mainly through the wage premium or penalty associated with these transitions.

These findings raise the important question of whether female workers, whose job changing rates are higher than those of male workers, are even more affected by changes in the relative strength of the U.S. dollar. We address this question in the current paper. We explore the effects of exchange rate movements on the wage growth of women, using an empirical specification that controls for differential wage growth by skill level. We estimate the impact of exchange rate shifts on: (1) the wages of women who remain with their same jobs, (2) the wages of women who change jobs, and (3) the frequencies of job-changing.

We find that women, like men, experience relatively more of their wage response to exchange rates at the time of job transitions. A 10 percent depreciation of the dollar, for example, is estimated to raise women's wages by roughly 1 percent. However, for women who have changed jobs the estimated wage increase is over 2 percent, while for women who stay on their jobs the estimated wage increase is 0.75 percent. For both genders, the strongest effects are observed among the lesser educated workers.

I. Theoretical and Conceptual Approach

A dynamic labor demand model motivates the linkage between exchange rates and labor market outcomes (Campa and Goldberg, 2001). Dollar movements affect a producer's revenues through both domestic and foreign market sales. Revenues arising from domestic sales respond to exchange rates in relation to import penetration in an industry. Similarly, revenues arising from foreign sales respond to exchange rates in relation to industry export intensity. Industry market structure is also likely to be important, with industries characterized by greater competition (and lower price-over-cost markups) expected to have larger labor demand elasticities. A producer's costs may also be impacted by exchange rates in relation to the degree to which the producer uses imported inputs.

Two additional results come out of the firm's interest in minimizing the costs of adjusting its labor force in response to external shocks. First, labor demand for higher wage and more skilled workers, who are more costly to replace, is less sensitive to revenue and cost shocks. Second, contemporaneous labor demand effects are expected to be strongest when exchange rate changes are perceived to be more permanent. Taken together, exchange rate effects on workers may differ in relation to industry competitive environment, trade exposure, and worker characteristics.

An important objective of our analysis is to link together the traditional studies of the effects of exchange rates on overall industry wages (Revenga, 1992; Campa and Goldberg, 2001) and the recent work on the effects of exchange rates on job-changing (Gourinchas, 1999; Goldberg, Tracy and Aaronson, 1999). To accomplish this, we begin by viewing the average wage as an expected wage for a specific skill-type of worker. This expected wage, for worker i employed in industry j at time t, is a probability-weighted average of the (expected) wage from remaining on the job in this industry during period t and the (expected) alternative wage available to a type i worker upon leaving for another job:[1]

$$E\left(w_t^i \middle| j \text{ at } t^-\right) = \left(1 - P_t^{ij}\right) E\left(w_t^i \middle| j \text{ at } t^+\right) + P_t^{ij} E\left(w_t^i \middle| k \text{ at } t^+\right) , \qquad (1)$$

where P_t^{ij} denotes the probability that worker i starting in industry j changes jobs at the start of period t and receives the alternative wage. The notation "j at t^-" indicates that a worker is located in industry j just prior to time t, and "k at t^+" indicates the job location of the worker just after time t. This index k reflects a job change that either places the worker into another industry or into another job within the broad industry j.

Differentiating the expected wage with respect to an industry-specific real exchange rate delineates the channels through which exchange rates influence the expected wages of workers:

$$\frac{\partial w_t^i / \partial rer_t^j}{w_t^i / rer_t^j} \Big| j \text{ at } t^- = \frac{\partial w_t^{ij} / \partial rer_t^j}{w_t^{ij} / rer_t^j} + P_t^{ij} \left(\frac{\partial \left(w_t^{ik} - w_t^{ij}\right) / \partial rer_t^j}{\left(w_t^{ik} - w_t^{ij}\right) / rer_t^j} \right) + \frac{\partial P_t^{ij} / \partial rer_t^j}{P_t^{ij} / rer_t^j} \left(w_t^{ik} - w_t^{ij}\right) \qquad (2)$$

As shown in equation (2), the overall wage elasticity reflects three mechanisms for wage adjustments. First, there is on-the-job wage adjustment, so that exchange rate fluctuations can affect a worker's wage in the absence of any employment transition. Second, given a normal frequency of job changing, the wage premium or penalty for job changes may be responsive to exchange rate movements. Third, given the normal wage premium or penalty for job-changing, exchange rate movements may induce a change in the rate of job-changing.

II. The Data

Our analysis uses the March Current Population Surveys (CPS) spanning the 1977 through 1998 survey years. Sample inclusion criteria restrict our population to civilian women between age 18 and 63, who are in private sector employment outside of Agriculture, Forestry, Fisheries and Mining.[2] At the level of the individual, the information we collect includes industry of employment, earnings, spells of unemployment and personal characteristics.[3]

The interview structure of the CPS allows us to potentially match half of the respondents to a March survey to the following year's March survey. However, a woman can be matched across adjacent March surveys only if she does not relocate during the intervening year. Our analysis starts with the "unmatched" pooled cross-sectional data consisting of about 337,000 workers,[4] of which we match 61 percent.[5]

Who is a "Job-Changer?" For a woman who reports positive earnings for a year, we classify her as a job-changer if she responded that she had more than one primary employer in the year, if she reported at least one spell of unemployment (not including temporary layoffs), if she worked less than 39 weeks during the year, or if she reported having worked in different industries.

In our unmatched sample of women with reported earnings, the job-changing rate is 0.40, implying that female workers on average change jobs nearly every 2.5 years. In any industry grouping, we observe the highest job-changing rates among the least educated workers. Workers in lower price-over-cost markup industries have higher job-changing rates, for any given level of educational attainment, compared with workers in other industries. These job-changing rates for women generally are higher than those observed for men with similar educational attainment and industry affiliations.

Our empirical analysis uses the permanent component of industry-specific real exchange rates. For a given industry and year, our exchange rate measure is constructed as the weighted average of the bilateral real exchange rates of U.S. trading partners, where the weights are the shares of each partner country in U.S. exports and imports for that industry. We compute the permanent component of the industry-specific series using a Beveridge-Nelson decomposition.[6]

III. Exchange Rates and Women's Expected Wages

A. Methodology. The wage response decomposition, described in equation (2), assumes workers of a common skill type. Our empirical specification isolates specific skill types by controlling for observed and unobserved heterogeneity in worker quality. We assume that the (log) wage for worker *i*, employed in industry *j*, and living in region *r* in period *t* is:

$$w_t^{ijr} = Z_t^i b + V_t^{jr} g + Y_t d + m_1^i + m_1^j + m_1^r + m_2^i t + m_2^j t + e_t^i \quad . \tag{3}$$

The vector Z_t^i controls for worker characteristics such as age, education, race, and marital status. Education is allowed to affect both the level and growth rate of wages. The vector V_t^{jr} contains industry-specific exchange rates as well as a Topel (1986)-type measure of local relative demand shocks.[7] Aggregate business cycles are captured in Y_t, which contains (log) real domestic GDP. Individual, industry and region-specific fixed effects as well as individual and industry-specific time trends are allowed, denoted by $m_1^i, m_1^j, m_1^r, m_2^i \cdot t$ and $m_2^j \cdot t$ respectively.

Since aggregate industry real wages, industry-specific real exchange rates and real GDP tend to display unit roots, using specification (3) to estimate the wage elasticities in the decomposition given by equation (2) could be problematic. To deal with this issue, we take advantage of the panel structure of our matched-CPS data and we first-difference the data across adjacent years. This yields the estimating equation:

$$\Delta w_t^{ijr} = \Delta Z_t^i b + \Delta V_t^{jr} g + \Delta Y_t d + m_2^i + m_2^j + \Delta e_t^i \tag{4}$$

An individual's wage growth is a function of her age, education, and changes in marital status. Industry-specific real exchange rates and aggregate real GDP variables now appear as growth rates. All of the fixed effects in our error structure drop out with the exception of the time trend terms which generate individual and industry fixed-effects in the wage growth specification. In estimation we control for industry fixed effects, but the individual fixed effects remain part of the error structure.

Our matched-CPS sample is not a random subsample of women workers from the March CPS surveys. To be matched across consecutive March surveys, a woman must remain in the same house over the intervening year. If the entire household moves or if the woman leaves the household, she can not be matched. Unlike panel data sets, the CPS makes no attempt to track down individuals who move. So, we must account for this sample selection issue when using the matched-CPS data to examine wage growth and job-changing rates. Another restriction imposed on our estimation sample is that a woman must have reported earnings. If a woman is out of the labor force for the entire previous year, we can not compute her wage rate.[8] We have to both match a woman across surveys and observe her earnings in order to compute her wage growth.

The sample selection issues implied by these restrictions are controlled for through inclusion of appropriate Mills' ratio terms. The first Mills ratio corrects for sample selection effects arising from matching across surveys, while a second Mills ratio corrects for sample selection effects arising from requiring a woman to have reported earnings. Given the lack of correlation between the residuals of the survey match and reported earnings specifications, we calculate these two Mills ratios using estimates from univariate Probit models.[9]

The Probit results show that, as expected, married women with children who own their own home are much more likely to be matched across surveys. The probability of being matched increases with age, and is higher for nonwhites and for individuals who have at least a high school education. The probability of having reported earnings increases with age and education, and is higher for whites and for women who rent their home. Married women with young children are less likely to have reported earnings. Only the selection based on reported earnings appears to be important in the wage growth specifications, so we drop the survey match Mills ratio. The coefficient on the reported earnings Mills ratio implies that individuals, who for unobserved reasons are more likely to have reported earnings, are also more likely to experience (again for unobserved reasons) faster wage growth.

The decomposition in equation (2) also requires that we estimate the impact of the exchange rate on the probability that a worker makes a job transition. We do this using Probit models where the control variables include all of the variables described in the wage growth specification. We also include an indicator for whether the woman changed jobs in the prior year to further control for heterogeneity in job-changing propensities across workers. We again control for sample selection in this estimation.

B. Differential Wage Responses for Job-Changers and Job-Stayers. The average wage elasticity for women with respect to an exchange rate depreciation is positive at 0.11 (Table 1, first column), and slightly larger than what we have observed for men.[10] This average masks a large disparity in elasticities between women who remain in their jobs at 0.08, and those who change jobs in the period following the exchange rate change at 0.20.

For job-changers, the point estimates of the wage elasticities can be quite large. Women with less educational attainment, especially within non-manufacturing, have the most responsive wages. Within manufacturing, the wage responses are largest in the lower price-over-cost markup industries (results not shown). This parallels our earlier findings for men that workers in industries with the lowest profit margins face the largest impacts from exchange rate fluctuations. The higher overall elasticity for women as compared to men primarily reflects the fact that the wage responsiveness for job-stayers is somewhat higher for women than men.

C. **The Mechanisms For Expected Wage Effects of Exchange Rates**. Table 2 presents the decomposition of the overall wage elasticities into the three channels outlined in equation (2), with the superscripts *s* and *c* denoting job-stayers and job-changers, respectively. As we observed for men, the expected wage effects of exchange rates at times of job transitions operate primarily through the direct wage consequences of these transitions ($P \cdot \delta [\Delta w^C - \Delta w^S]/\delta\Delta\, rer$), rather than through the incidence of these transitions ($([\Delta w^C - \Delta w^S] \cdot \delta P/\delta\Delta rer)$). On average, changing jobs is associated with a 28 percent lower wage growth. This wage penalty for job-changing can decline enormously in response to a positive labor demand shock, such as a dollar depreciation. The combination of already high job-changing rates (captured by P) with the highly sensitive relative wage growth between job-stayers and job-changers makes dollar fluctuations potentially a significant concern for women. Exchange-rate induced changes in job changing rates (Table 2, last column) account for little of the overall wage effects for women.

IV. Concluding Remarks

Our findings show that women, like men, experience most of the expected wage response to dollar fluctuations at times of job-transitions, rather than when they remain with the same employer. In this context, dollar depreciation periods — generally viewed as providing positive labor demand shocks — reduce the penalties that often are associated with a job change. Since women have higher job changing rates than their male counterparts, these findings suggest that the average female worker has more sensitive wages than her male counterpart. Within the population there is diversity in these effects, with the least educated women having both the highest job transition rates and the largest wage response to exchange rates at these transitions.

Table 1	Wage Elasticities for Women, by education & industry groups and job changing status		
Industry / Education Group (sample size: overall = job-stayers + job-changers)	Overall	Job-stayers	Job-changers
All Workers (94,209 = 68,273 + 25,936)	0.11** (0.04)	0.07 (0.08)	0.20* (0.11)
High school dropouts (13,250 = 8,912 + 4,338)	0.06 (0.08)	–0.06 (0.11)	0.31 (0.20)
High school graduates (44,898 = 32,816 + 12,082)	0.16** (0.05)	0.10* (0.06)	0.33** (0.16)
Some college + (36,061 = 26,545 + 9,516)	0.05 (0.13)	0.09 (0.11)	–0.02 (0.24)
Manufacturing (20,077 = 5,010 + 5,067)	0.10* (0.06)	0.03 (0.08)	0.15 (0.21)
High school dropouts (5,112 = 3,555 + 1,557)	–0.04 (0.15)	–0.07 (0.14)	–0.12 (0.38)
High school graduates (10,187 = 7,700 + 2,487)	0.15 (0.10)	0.05 (0.07)	0.28 (0.28)
Some college + (4,778 = 3,755 + 1,023)	0.19 (0.26)	0.12 (0.25)	0.43 (0.66)
Non-manufacturing (74,132 = 53,263 + 20,869)	0.11* (0.05)	0.08 (0.08)	0.21 (0.12)
High school dropouts (8,138 = 5,357 + 2,781)	0.12 (0.16)	–0.06 (0.19)	0.55** (0.20)
High school graduates (34,711 = 25,116 + 9,595)	0.17** (0.06)	0.11* (0.07)	0.34 (0.23)
Some college + (31,283 = 22,790 + 8,493)	0.03 (0.13)	0.09 (0.12)	–0.07 (0.25)

Notes: ** significant at the 5% level, * significant at the 10% level. Reported coefficients are elasticities with respect to a dollar depreciation. Each elasticity is estimated from a separate regression. See text for control variables. Standard errors have been adjusted for any non-independence of observations within years.

Table 2	Decomposition of Expected Wage Elasticities for Female Workers			
Sample	Implied Overall Elasticity	Job-Stayers	Incremental Effects Associated with Job-Change	
		$\frac{\partial \Delta w^s}{\partial \Delta rer}$	$P \bullet \frac{\partial \left(\Delta w^c - \Delta w^s \right)}{\partial \Delta rer}$	$\left(\Delta w^c - \Delta w^s \right) \bullet \frac{\partial P}{\partial \Delta rer}$
Private nonagricultural				
All Workers	0.11	0.05 (0.08)	0.27 • 0.21 = 0.06 (0.14)	–0.28** • 0.00= –0.00 (0.01) (0.02)
High school dropouts	0.04	–0.10 (0.12)	0.33 • 0.47* = 0.15 (0.28)	–0.35** • 0.05 = –0.01 (0.01) (0.08)
High school graduates	0.16	0.08 (0.07)	0.27 • 0.31* = 0.08 (0.18)	–0.29** • –0.02 = 0.01 (0.01) (0.13)
Some college +	0.05	0.07 (0.12)	0.26 • –0.04 = –0.01 (0.21)	–0.25** • 0.01 = –0.00 (0.01) (0.03)
Manufacturing				
All Workers	0.10	–0.01 (0.09)	0.25 • 0.29 = 0.07 (0.23)	–0.32** • –0.10* = 0.03 (0.01) (0.05)
High school dropouts	–0.09	–0.14 (0.14)	0.30 • 0.15 = 0.04 (0.38)	–0.36** • –0.01 = 0.00 (0.02) (0.12)
High school graduates	0.15	0.01 (0.11)	0.24 • 0.42 = 0.10 (0.33)	–0.30** • –0.11 = 0.03 (0.02) (0.07)
Some college +	0.17	0.08 (0.25)	0.21 • 0.35 = 0.07 (0.60)	–0.32** • –0.05 = 0.01 (0.03) (0.10)
Non-Manufacturing				
All Workers	0.11	0.06 (0.08)	0.28 • 0.18 = 0.05 (0.14)	–0.27** • 0.01 = –0.00 (0.01) (0.03)
High school dropouts	0.11	–0.08 (0.18)	0.34 • 0.65** = 0.22 (0.28)	–0.34** • 0.09 = –0.03 (0.01) (0.10)
High school graduates	0.17	0.09 (0.08)	0.27 • 0.27 = 0.07 (0.26)	–0.29** • 0.00 = –0.00 (0.01) (0.03)
Some college +	0.03	0.06 (0.13)	0.27 • –0.08 = –0.02 (0.23)	–0.24** • 0.03 = –0.01 (0.01) (0.04)

Notes: ** significant at the 5% level. * significant at the 10% level. The decomposition is calculated with respect to a dollar depreciation. For each row, the job-stayer elasticity, the average job-stayer versus job-changer wage growth differential, and the wage growth differential elasticity are estimated from a single regression. The implied overall elasticity is the sum of the other three terms in the row. Standard errors have been adjusted for any non-independence of observations within years.

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[1] The expectations operator is omitted for notational convenience.

[2] The sample is further limited to workers who were not in school and who are not primarily self-employed. We symmetrically trimmed the top and bottom 1 ½ percent of workers based on income to eliminate top-coding of the income variable.

[3] The March t+1 survey for any individual contains current period variables and year t information. Thus, given survey years 1977 through 1998, we have wages for 1976 through 1997.

[4] This is the sample of workers in their first March interview who satisfy our inclusion criteria and who have no missing values for our variables of interest (excluding earnings).

[5] Due to survey limitations, we are unable to match workers for the survey years 1985/86 and 1995/96.

[6] See Campa and Goldberg (forthcoming 2001) for a more extensive discussion of these series.

[7] See Goldberg and Tracy (2001). The constructed local relative demand shock measures a state employment shock as a deviation from the national employment shock.

[8] We treat allocated earnings as if they are missing.

[9] Three variables are included in the Probit specifications which are excluded from our wage growth specification. These include two variables for the presence of children in the household and an indicator for whether the household head owns or rents the home. In addition, the Probit specifications includes year effects, and marital status enters in level form (marital status enters the wage growth specifications in change form).

[10] We limit the sample to workers who would be interviewed in the following March survey if they did not move. Our methodology constrains the effects of dollar appreciations and

depreciations to be symmetric. Separate results for the higher and lower markup manufacturing industries are not reported.