Exchange Rates and Wages

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Abstract

Understanding the effects of exchange rate fluctuations across the population is important for increasingly globalized economies. Previous studies using industry aggregate data have found that industry wages are significantly more responsive than industry employment to exchange rate changes. We offer an explanation for this paradoxical finding. Using Current Population Survey data for 1976 through 1998, we document that the main mechanism for exchange rate effects on wages occurs through job turnover and the strong consequences this has for the wages of workers undergoing such job transitions. By contrast, workers who remain with the same employer experience little if any wage impacts from exchange rate shocks. In addition, we find that the least educated workers –who also have the most frequent job changes – shoulder the largest adjustments to exchange rates.

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

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I. Introduction

As large swings in the exchange rates become a regular part of the economic landscape, their implications for workers need to be established. An evolving literature suggests that the wage effects of dollar movements can be significant, but there remain broad gaps in our understanding of the implications for different groups within the workforce and of the mechanisms for these effects. In this paper, we confront these issues using a rich source of data on labor market outcomes for individual workers employed in U.S. manufacturing and non-manufacturing industries. We find that exchange rate movements lead to sizable wage changes for some workers, especially the less educated portion of the U.S. workforce. We show that the incidence of these wage implications is primarily at times of worker job transitions.

Our results unify and significantly expand upon the insights from prior analyses. Using industry-level aggregates, prior studies have identified a paradox, in that exchange rate movements appear to have little effect on jobs, but sometimes sizable effects on wages. Campa and Goldberg (forthcoming) study the full range of U.S. manufacturing industries for 1972 through 1995 using a broad set of margins for labor market response (wages, hours, jobs, overtime wages and overtime hours) and document large industry wage responsiveness but considerably smaller and often insignificant employment elasticities in response to dollar movements.[1] Industries were heterogeneous in these effects: those with lower price-over-cost markups — interpreted as facing stiffer competitive conditions — have smaller wage responsiveness and larger employment responsiveness than higher-markup industries, while industries with smaller proportions of skilled-labor have less wage and more employment responsiveness. Higher price-over-cost margin industries, which also on average are more skilled labor-intensive in production, rely more intensively on adjustments to overtime work and pay.

This finding of high wage responsiveness but relatively low employment responsiveness is paradoxical given the expectation that wages will be sticky and employment more flexible, instead of the other way around. It raises interesting questions about the informativeness of industry-level data and about the potential mechanisms for exchange rate effects.

One caution is that industry-level studies may provide misleading insights into the implications of currency fluctuations for individual workers. For example, substantial changes in gross employment (i.e. job churning or turnover) may be largely missed in the aggregated data.

Observed employment responsiveness captures only net employment changes and only among those workers who move across but not within broadly defined industries. The issue of aggregation comes to the fore in recent work on exchange-rate-induced churning of workers.[2]

Gourinchas (1998) shows that, despite the relatively small implications of exchange rates on net employment within industries, dollar depreciations (appreciations) lead to a significant and large chill in (stimuli to) gross employment changes. Labor reallocations within industries decline as gross job creation and gross job destruction both fall significantly along with the depreciation of the dollar. Goldberg, Tracy and Aaronson (1999) reach similar conclusions based on analysis of a panel of workers drawn from the *Current Population Survey* (CPS). They find that dollar depreciations reduce industry switching by male workers in manufacturing industries. These effects are strongest in non-durables industries, especially those facing higher import penetration. For male workers in non-manufacturing industries, the signs and significance of exchange rate effects are more mixed.

Wage data are subject to a distinct set of aggregation issues: time series data indicate the average wages of groups of individuals who are employed at different dates, without regard for the potential changes in the skill composition of these worker groups. For example, if dollar appreciations induce relatively more turnover of low-skilled workers in an industry, then this creates a positive composition bias to the observed change in the industry average wage.[3]

In the present paper we systematically address a number of key questions about the incidence of exchange rate effects on labor markets. First we document which types of workers are particularly sensitive to exchange rates, focusing on individuals delineated by both industry of employment and by their educational attainment. We consider the implication of dollar movements for the wages of these individuals and for the frequency with which they move between jobs. We also focus particular attention on the mechanism through which wages respond, with the goal of shedding light on the aforementioned paradox. Specifically, do workers have to change jobs (within or across industries) in order to have their wages influenced by dollar fluctuations, or are these implications felt by workers who remain with the same

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1 See Branson and Love (1988) and Revenga (1992) for earlier industry studies.

2 This type of churning has been shown to be sensitive to other macroeconomic shocks, including oil price fluctuations. Davis and Haltiwanger (1999) show that job destruction rates are more sensitive to oil price shocks than are job creation rates in every sector with the clear exception of small, young plants.

3 This same issue comes up in trying to measure the cyclicality of real wages. See Keane and Prasad (1993) and Solon and Barsky (1994).

employer? A related issue is whether dollar fluctuations have any economically important influences on the probability of job-changing by different types of individuals, particularly since job-changing is usually associated with wage adjustments.

We tackle these themes using micro-labor market data for male respondents to the CPS for the period 1976 through 1998. This data presents us with a key advantage over previous work, since we have the ability to control for the characteristics of individual workers. We match workers across adjacent March interviews, creating a series of two-year panels on wages and employment status of individuals. This allows us to explore the effects of exchange rate movements on the wage growth of these workers, using an empirical specification that controls for differential wage growth by skill level. We trace the mechanisms for wage responsiveness by linking the effects to: (1) workers who remain with their same jobs, (2) workers who switch jobs, and (3) exchange-rate induced changes in the frequency of job-switching by particular classes of workers. We document the importance of these mechanisms for workers sorted by major industry and educational attainment. While the overall average wage elasticity with respect to the exchange rate is around 0.08, we find that the elasticity for low-skilled job changers to be around 0.5. For most skill groups and industries, the impact of dollar fluctuations on wages is concentrated among those workers who change jobs.

The paper is organized as follows. Section II provides the theoretical underpinnings for the wage and exchange rate interactions to be examined empirically. Section III describes our main data, and details the criteria and methods used in choosing the sample population, constructing the job-changing frequencies, and calculating industry-specific exchange rates. Section IV presents our empirical methodology and results. Section V offers concluding remarks.

II. Theoretical and Conceptual Approach

To motivate the linkage between exchange rates and labor market outcomes, our starting point is the theoretical structure on labor demand derived in Campa and Goldberg (forthcoming) and Goldberg and Tracy (2000). The sensitivity of labor demand to exchange rates arises primarily through the impacts of exchange rates on producer profits. Dollar movements can affect a producer's revenues through impacts on domestic and foreign market sales. A producer's factor costs may also vary with respect to exchange rates.

In this setup, domestic production uses three inputs: imported components, domestic components, and domestic labor. Labor inputs are costly to adjust due to training, hiring and firing fees that are assumed to rise in relation to a worker's wage. The model generates a dynamic labor demand function wherein the effects of exchange rates depend on the level and form of industry trade orientation, the competitive structure of the industry, and the intensity of labor usage in production. Due to the adjustment costs incurred when employment levels are changed, the willingness of producers to adjust employment is higher when exchange rate fluctuations are perceived to be more permanent.

The solution to the producer's dynamic labor demand problem yields an elasticity of labor demand *L* with respect to the industry-specific real exchange rate, rer^j, proportional to:[4]

$$\frac{\partial L_t^j}{\partial rer_t^j} \Big/ \frac{L_t^j}{rer_t^j} \propto \frac{p_t^j}{\beta^j}\left(\left(1+\eta^{j^{-1}}\right)\kappa^j M_t^j + \left(1+\eta^{j\,*^{-1}}\right) X_t^j - \left(\partial Q^j / \partial Z^*\right)^{-1} \alpha_t^j\right) \ , \qquad (1)$$

where rer_t^j is defined as domestic currency per unit of foreign exchange for any industry *j*, p_t^j is the unit price of output, Q^j is the production function for industry *j*, *Z** is the quantity of imported inputs, η^j and η^{j*} are domestic and foreign elasticities of demand, β^j is the share of labor used in production, $\kappa^j M^j$ indicates that the sensitivity of home market prices with respect to exchange rates is proportional to the import penetration of domestic markets (*M*), with κ as the index of proportionality, X^j is the export orientation of the industry, and α^j is the production share of imported inputs. The degree of proportionality arises because the elasticity is increasing in the perceived permanence of the exchange rate movement and smaller for the higher-wage and more skilled workers for whom replacement is more costly.

Equation (1) emphasizes the three main transmission channels through which exchange rates influence labor demand. First, the greater the import penetration of domestic markets (M_t^j), the greater the revenue implications of exchange rates: this result occurs because of price pressures that can undermine a local producer's competitiveness in home markets. Second, higher export-orientation (X_t^j) raises the sensitivity of labor demand to exchange rates by making a larger fraction of sales revenues more directly linked to currency fluctuations. Third, greater reliance on imported components (α_t^j) expands the potential production cost ramifications of exchange rates.

[4] Solution details are in Campa and Goldberg (forthcoming).

Since the cost implications are opposite in sign to the revenue implications, reliance on imported inputs can potentially offset the adverse revenue consequences of a dollar appreciation.

Equation (1) also highlights particular industry features that magnify or reduce the induced degree of industry labor demand shifts. More labor-intensive production (i.e. higher β^{j}) reduces the overall sensitivity of labor demand to exchange rates. Industries characterized by greater competition among firms (with lower demand elasticities η^{j} or η^{j*}) are expected to have larger labor demand elasticities. In addition to exchange rate terms, other variables that influence labor include domestic aggregate and local demand conditions, world demand conditions, and the prices of other factor inputs.

This derived labor demand expression motivates the form with which exchange rates should enter regression specifications, and suggests sample stratifications that are likely to provide useful empirical insights. One justifiable sample split is along industry lines, since labor demand responsiveness will be higher for industries facing stiffer competitive conditions. Another appropriate sample division is based on differences in worker skill levels, especially to the extent that these are associated with the implied costs of labor market adjustments. A third potential stratification is via export orientation, import penetration, and reliance on imported inputs by industries. We opt against using this for a sample split, and instead test for the evolving effect of exchange rates by interacting exchange rate fluctuations with measures of industry exposure to external trade.

An approximate log-linearization of optimal labor demand by industry *j* located in local labor market *r* is given by:

$$L_t^{jr} = c_0^j + c_1 y_t + c_2 y_t^r + c_3 y_t^* + (c_{41}\chi_t^j + c_{42} M_t^j + c_{43}\alpha_t^j)\cdot rer_t^j + c_5 w_t^j + c_6 s_t + c_7 L_{t-1}^j \quad (2)$$

where all variables other than χ_t^j, M_t^j and α_t^j are defined in logarithms. Shocks that are common to all industries include aggregate demand conditions y_t, local demand conditions y_t^r, world demand conditions y_t^*, and the cost of capital s_t. The industry wage is given by w^j. Lagged employment terms are introduced only when there is costly adjustment and some perceived persistence in exchange rate shocks.

We adopt a simple form for labor supply to an industry. We assume that labor supply is a function both of the level of industry wages and the industry wage relative to the alternative wage in the locality, w_t^{ar}. The alternative wages in a locality is positively related to current local demand conditions as in Topel (1986).

A reduced form representation for overall labor supply to an industry *j* located in local labor market *r* is given by:

$$L_t^{jr} = a_0^j + a_1 w_t^j + a_2 \left(w_t^j - w_t^{ar} \right), \tag{3}$$

which allows for industry-specific fixed effects.

The solution to equations (2) and (3) are equilibrium outcomes for employment and wages for any industry *j* given by:

$$w_t^j = \omega_1^j + \omega_2 y_t + \omega_3 y_t^r + \omega_4 y_t^* + \omega_5 w_t^{ar} + (\omega_{6,1}^j \chi_t^j + \omega_{6,2}^j M_t^j + \omega_{6,3}^j \alpha_t^j) rer_t^j + \omega_7 s_t + \omega_8 L_{t-1}^j \tag{4a}$$

$$L_t^j = \lambda_1^j + \lambda_2 y_t + \lambda_3 y_t^r + \lambda_4 y_t^* + \lambda_5 w_t^{ar} + (\lambda_{6,1} \chi_t^j + \lambda_{6,2} M_t^j + \lambda_{6,3} \alpha_t^j) rer_t^j + \lambda_7 s_t + \lambda_8 L_{t-1}^j \tag{4b}$$

These equations are similar to the industry-level specifications previously estimated in the literature. However, our intent is to use wage data at the level of individual workers. We assume that the intercept in (4a) is a function of observed individual characteristics, Z^i, such as age, education, race and marital status. Education is allowed to affect both the level and growth rate of wages. Then, for ease of exposition, we collapse the time-varying terms from (4a) into vectors V_t^{jr} and Y_t. V_t^{jr} contains industry-specific exchange rates and a measure of local labor market relative demand shocks. Y_t captures aggregate business cycles, represented by the path of real domestic GDP.[5] Finally, the error term is modeled using an error components structure which allows for individual, industry and region-specific error components as well as individual and industry-specific time trends, denoted by $\mu_1^i, \mu_1^j, \mu_1^r, \mu_2^i \cdot t$ and $\mu_2^j \cdot t$ respectively. Accordingly, we rewrite (4a) so that the (log) wage for worker *i*, employed in industry *j*, living in region *r* in period *t* is given by:

$$w_t^{ijr} = Z_t^i \beta + V_t^{jr} \gamma + Y_t \delta + \mu_1^i + \mu_1^j + \mu_1^r + \mu_2^i t + \mu_2^j t + \varepsilon_t^i \quad . \tag{5}$$

Since aggregate industry real wages, industry-specific real exchange rates and real GDP tend to display unit roots, using specification (5) to estimate the wage elasticities would be problematic. To deal with this issue, we take advantage of the panel structure of our matched-CPS data and first-difference the data across adjacent years. This yields the estimating equation:

$$\Delta w_t^{ijr} = \Delta Z_t^i \beta + \Delta V_t^{jr} \gamma + \Delta Y_t \delta + \mu_2^i + \mu_2^j + \Delta \varepsilon_t^i \qquad (6)$$

An individual's wage growth is a function of his 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 error components drop out with the exception of the individual and industry-specific time trends, which generate individual and industry-specific error components in the wage growth rate.[6] In the estimation we control for the remaining industry-specific error components, μ_2^j, using industry fixed-effects, but the individual-specific error components, μ_2^i, remain part of the composite error term.

Thusfar, the estimation approach provides information about the average wage consequences of movements in exchange rates and other forcing variables. Since we also are interested in the mechanisms for the exchange rate effects we can decompose the expected wages of an individual of skill type *i*, working in industry *j* at time *t*. This expected wage is a probability-weighted average of the (expected) wage from remaining on the job in this industry during all of period *t* and the (expected) alternative wage available to a type *i* worker upon leaving for another job:

$$E\left(w_t^i \middle| j \text{ at } t^-\right) = \left(1 - P_t^{ij}\right) E\left(w_t^i \middle| j \text{ at } t^+\right) + P_t^{ij} E\left(w_t^i \middle| k \text{ at } t^+\right) \qquad (7)$$

where "*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* indicates a job change that can place the worker into other industries or permits change within the broad industry *j*. P_t^{ij} denotes the probability that this worker *i* starting in industry *j* changes jobs at the start of period *t* and receives the alternative wage.

[5] We do not control for world GDP, y^*, the cost of capital, *s*, or lagged industry employment, L^j.

[6] In particular, first-differencing further controls for unobserved worker quality since μ^i_1 drops out of the

Differentiating equation (7) with respect to industry-specific real exchange rates, we derive the set of channels through which exchange rates influence the expected wages of individual workers. Specifically, for workers of type i, the elasticity of expected wages with respect to an exchange rate movement at the beginning of period t is

$$\frac{\partial w_t^i / \partial rer_t^j}{w_t^i / rer_t^j}\Big| j \text{ at } t^- = \frac{\partial w_t^{ij} / \partial rer_t^j}{w_t^{ij} / rer_t^j} \quad + P_t^{ij}\left(\frac{\partial\left(w_t^{ik} - w_t^{ij}\right) / \partial rer_t^j}{\left(w_t^{ik} - w_t^{ij}\right) / rer_t^j}\right) + \frac{\partial P_t^{ij} / \partial rer_t^j}{P_t^{ij} / rer_t^j}\left(w_t^{ik} - w_t^{ij}\right) \quad (8)$$

where the expectations operator is omitted for notational convenience.

Equation (8) clearly depicts the three mechanisms for wage adjustments from the vantage point of an individual worker. First, there can be on-the-job wage adjustment, so that exchange rate fluctuations affect the worker in the absence of employment transitions. Second, given a normal frequency of job churning, the associated wage premium or penalty 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 churning. Empirically, we jointly estimate $\partial w_t^{ij} / \partial rer_t^j$ and $\partial\left(w_t^{ik} - w_t^{ij}\right) / \partial rer_t^j$ by introducing into specification (6) appropriate dummy variables to distinguish between job-stayers and job-changers. We also specify a Probit model on job-changing for different types of works to estimate $\partial P_t^{ij} / \partial rer_t^j$.

III. The Data

A. Matched CPS Data. To explore the extent of wage effects of exchange rates and the mechanisms for these effects, we use micro data from the March CPS spanning the 1977 through 1998 survey years, which provide wage information for 1976 through 1997. Sample inclusion criteria restrict our population to civilian men between age 18 and 63, in private sector employment outside of Agriculture, Forestry, Fisheries and Mining.[7]

We collect a range of other information for each individual so that we can determine

specification.

[7] The sample is further limited to workers who were not in school, who were not primarily self-employed and who had positive weeks worked and earnings in both years. We symmetrically trimmed the top and bottom 1 ½ percent of workers based on income. This effectively eliminates the problem of top-coding of earnings data. We also

whether the same individual has been surveyed at the same location across two consecutive survey years. For these cases, we can match the individual's two surveys and compute the individual's wage growth over the two years. However, we also need to examine the broader "unmatched" sample to determine whether this matching process leads to sample selection biases (and ultimately correct for such biases). The unmatched sample consists of 204,690 individuals with reported earnings who can potentially be matched; the "matched" sample consists of 114,404 individuals. Due to survey limitations, we are unable to match workers for the survey years 1985/86 and 1995/96.

The unmatched and matched samples of workers have similar characteristics (Appendix Table A1). There are also notable differences, with the most obvious being the lower job-changing rates in the matched sample. This is expected, since the matched sample is comprised of individuals identified at the same location at two consecutive survey dates. Matched individuals are more heavily concentrated in manufacturing industries, less likely to be high school dropouts, more likely to be married and homeowners, and are slightly older than their unmatched counterparts.

The matched sample is well-distributed across different educational and industry groupings. Within manufacturing, we further differentiate among industries according to competitive structure, as proxied by their price-over-cost markup histories.[8] Our interpretation is that producers in the Lower (price-over-cost) Markup group have less pricing power (and more elastic product demand) than those producers in Higher-Markup industries. All else equal, the wage effects of exchange rates should be higher in the lower markup industries, as discussed in Section II. Of course, these effects can also differ across skill classes of workers, which motivates our distinction by the educational attainment of workers. Lower-markup industries employ the greatest proportion of low education workers, with more than 70 percent of the workforce as either high-school dropouts or high school graduates. Workers in high markup manufacturing industries have similar educational profiles to workers in the non-manufacturing

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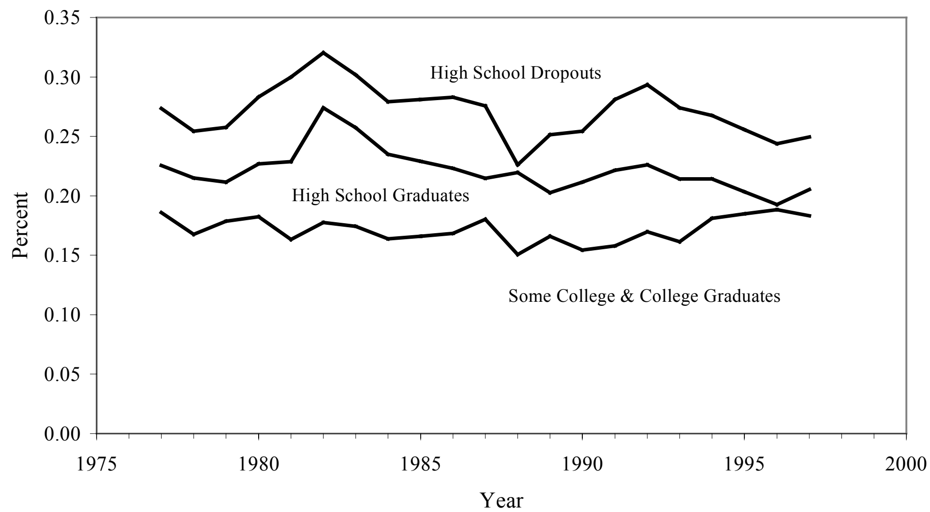
eliminate observations with allocated or missing values for the variables we use in the analysis.

[8] Domowitz, Hubbard and Peterson (1986) show that average markups for industries are highly correlated with industry competitive structure. The markups are calculated using a value-added measure: (Value Added plus Cost of Materials)/(Payroll plus cost of materials), with data from the Annual Survey of Manufactures from the Bureau of the Census. Within manufacturing, the Lower Markup industries are: primary metal products, fabricated metal products, transportation equipment, food and kindred products, textile mill products, apparel and mill products, lumber and wood products, furniture and fixtures, paper and allied products, petroleum and coal products, and leather and leather products.

industries, both with higher shares of college-educated employees.

B. Job-Changing Measures. We define a worker as a "job changer" in year *t* if he responded that he had more than one primary employer in that year or if he reported at least one spell of unemployment (not including temporary layoffs). A worker also is classified as a job changer if he worked less than 39 weeks during year *t* or reported having worked in different industries over the consecutive survey dates.

Figure 1. Job Changing Rates: by Worker Educational Attainment



Notes: Turnover rates could not be calculated from the CPS data for 1985 and 1995. We filled in these missing turnover rates using linear interpolation.

Job-changers, on average, have different characteristics than "job-stayers" (Appendix Table A1). Job-changers are younger, have lower levels of educational attainment, are less likely to be married, and are more likely to have been employed in non-manufacturing industries. In the matched sample, annual job-changing rates are lowest for workers in high-markup industries (at 15 percent, or once every 6.7 years) and highest for workers in the non-manufacturing industries (at 23 percent or once every 4 years). There also are tremendous differences in job-changing rates in accordance with educational attainment. The least educated workers (without high school degrees) in non-manufacturing industries frequently change jobs (at 32 percent, about once every 3 years). For workers with at least some college degree the rate of job changing is much lower, averaging 19 percent (once every 5 years) in non-manufacturing industries, and 14 percent (once every 7 years) in manufacturing. As evident in Figure 1, job changing rates for less educated men are more sensitive to the business cycle.

C. Industry-Specific Real Exchange Rates. For a given industry and year, we construct industry-specific real exchange rates as a weighted average of the bilateral real exchange rates of U.S. trading partners. The weights vary across industries and over time, and are the shares of each partner country of the U.S. in industry exports or imports (trade data are from U.S. Department of Commerce; bilateral exchange rates are from the *International Financial Statistics*).[9]

As derived in Campa and Goldberg (forthcoming), exchange rate implications for labor outcomes should be strongest for the permanent component of the exchange rate. We compute this permanent component by applying a Beveridge-Nelson decomposition over quarterly data (see Appendix description). The resulting process is used for generating the annualized exchange rate observations that are used in our regression specifications.

D. Aggregate and Local Demand Variables. We control for the U.S. business cycle by including real GDP in the regressions. Since local labor market conditions are relevant for local outcomes, we follow Topel (1986) and compute a measure of, the local relative labor market demand conditions, y_t^r. For each state *r* of the United States, we regress the logarithm of private-sector nonagricultural employment in that state on a quadratic time trend. The residuals from these regressions, γ_t^r, measure the deviations from trend employment in state *r* at time *t*. Similarly, we regress the logarithm of national private sector non-agricultural employment on a quadratic time trend. The residuals from this regression, γ_t,control for the aggregate business cycle. The local relative demand shock in state *r* and in year *t* is defined as

$$y_t^r = \gamma_t^r - \gamma_t \quad , \tag{9}$$

[9] We used trade shares for bilateral transactions with 34 trading partners of the United States. Our methodology for the weighted export plus import share exchange rate parallels the approach recently implemented by the Federal Reserve Board of Governors. Our mapping of exchange rates to CPS industries is available upon request. These industry-specific exchange rates are more appropriate for our empirical study than a single aggregate exchange rate. Due to differences in international product destinations and sources of import competition, a single exchange rate measure for the United States is sometimes a poor measure of the relevant fluctuations in exchange values.

so that the local relative demand shock measures the period t state employment shock as a deviation from the national employment shock. Larger values represent relatively tighter conditions in that local labor market than for the country as a whole.

E. Industry trade exposure. Since the theoretical motivation suggests that exchange rate implications for workers may vary with industry trade orientation, we interact the real exchange rates with industry-specific export and import trade shares in some regression specifications. At each date, we have industry-specific values of exports and imports (Department of Commerce series). The industry export share is computed by dividing industry exports by industry shipments. Industry imports are deflated by industry shipments less exports to provide a measure of industry import penetration.

IV. Estimation and Results

The empirical analyses reported in this section provide estimates of expected wage elasticities with respect to exchange rates, where we focus on the importance of industry affiliation and educational attainment. We then decompose the channels for these wages effects, identifying whether they arise for workers who remain on the job, through the wage penalty for workers who change jobs, or through a change in the probability of job transitions. Before presenting these findings, Section A discusses our estimation methods, including our careful controls for potential sample selection issues.

A. Sample Selection. As is clear from the descriptive statistics given in Appendix Table A1, our matched-CPS sample is not a random subsample of male workers from the March CPS surveys. To be matched across consecutive March surveys, a worker must remain in the same house over the intervening year. If the entire household moves or if this worker leaves the household, he 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.

A second restriction imposed on our estimation sample is that a male worker must have reported earnings. If a man is out of the labor force for the entire previous year or if he is employed but refuses to answer the earnings question, then we can not compute the growth in his wages.

We have to both match a man across surveys and observe his wages in order to compute his wage growth. The first step toward controlling for these two sources of sample selection is to empirically model the processes for matching an individual across March surveys and for having reported earnings. Given that non-matches are primarily due to geographic mobility, our empirical specification borrows from the migration literature.[10] This literature shows that mobility is strongly tied to many individual characteristics such as age, education, and marital status. In addition, factors that increase the transaction costs of moving reduce mobility. These factors include owning a house and having children. We also use these same demographic factors to help control for whether a man has reported earnings.

We use the following methodology for controlling for sample selection biases. Let I_{1t}^{i*} denote an unobserved index of the desire by an individual i to stay in the same house during period t. We use the normalization that if this index exceeds zero, then the man does not move and we can match him across surveys. Let I_{1t}^{i} denote an indicator that takes a value of one if we match the individual, and zero if we can not match him. Similarly, let I_{2t}^{i*} and I_{2t}^{i} denote the unobserved and observed indices for whether a man has reported earnings. Let X^i denote a vector of demographic characteristics for individual i that affect his mobility and propensity to report earnings. Combining these indices with our earlier wage growth equation, we have the following joint specification where we have relabeled the composite residual in the wage growth equation for ease of exposition, i.e. $n_{3t}^{i} = m_2^i + m_2^j + \Delta e_t^i$.

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\begin{aligned}
&I_{1t}^{i*} = X_1^{i\prime} q_1 + n_{1t}^{i} \\
&I_{1t}^{i} = 1 \text{ if } I_{1t}^{i*} > 0, 0 \text{ otherwise} \\
&I_{2t}^{i*} = X_{2t}^{i} q_2 + n_{2t}^{i} \\
&I_{2t}^{i} = 1 \text{ if } I_{2t}^{i*} > 0, 0 \text{ otherwise} \\
&\Delta w_t^{ijr} = \Delta Z_t^i b + \Delta V_t^{jr} g + \Delta Y_t d + n_{3t}^{i} \\
&\begin{pmatrix} v_{1t}^{i} \\ n_{2t}^{i} \\ n_{3t}^{i} \end{pmatrix} \sim N(0, \Sigma) \text{ where } \Sigma = \begin{bmatrix} 1 & s_{12} & s_{13} \\ \bullet & 1 & s_{23} \\ \bullet & \bullet & 1 \end{bmatrix}
\end{aligned} \qquad (10)
$$

[10] See for example, Bartel (1979).

The expression for the expected wage growth, conditional on an individual being matched across surveys and having reported earnings, depends on whether there is any correlation between the residuals in the survey matching equation, I_{1t}^{i*}, and the reported earnings equation, I_{2t}^{i*}. We check the extent of this correlation, σ_{12}, by estimating a joint Probit model for whether we match an individual across March surveys, and whether the individual has reported earnings in the base year survey. We find that σ_{12} equals 0.13. Given the low value of σ_{12} we assume that the residuals in the match and reported earnings equations are independent. In this case, the expression for the expected wage growth conditional on an individual being matched across surveys and having reported earnings, becomes:

$$\begin{aligned} E(\Delta w_t^{ijr} \mid I_{1t}^i = 1, I_{2t}^i = 1) &= \Delta Z_t^i \beta + \Delta V_t^{jr} \gamma + \Delta Y_t \delta + E(v_{3t}^j \mid v_{1t}^j > -X_{1t}^i \theta_1) + E(v_{3t}^j \mid v_{2t}^j > -X_{2t}^i \theta_2) \\ &= \Delta Z_t^i \beta + \Delta V_t^{jr} \gamma + \Delta Y_t \delta + \sigma_{13} \frac{\phi(X_{1t}^i \theta_1)}{\Phi(X_{1t}^i \theta_1)} + \sigma_{23} \frac{\phi(X_{2t}^i \theta_2)}{\Phi(X_{2t}^i \theta_2)} \end{aligned} \qquad (11)$$

where φ( ) and Φ( ) are the standard normal density and cumulative density functions.

The first Mills ratio corrects for sample selection effects arising from matching across surveys, while the second Mills ratio corrects for sample selection effects arising from requiring a worker to have reported earnings. Given the lack of correlation between the match and reported earnings specifications, we calculate these two Mills ratios using θ's estimated from univariate Probit models.[11]

In addition to estimating wage elasticities with respect to the exchange rate, we estimate the impact of the exchange rate on the probability that a worker makes a job transition between the March surveys. Here we estimate Probit models where we again control for sample selection. The control variables in these Probits include all of the variables described in the wage growth specification in addition to an indicator for prior job mobility.

The Probit results used to calculate the two Mills ratios are summarized in Appendix Table A2.[12] As expected, married men with older children who own their own home are much

[11] Here we limit the sample to those workers who would be interviewed in the following March survey if they did not move.

[12] Three variables are included in these specifications which are excluded from our wage growth specification. These include two variables for the presence of children in the household, and a variable indicator for whether the

more likely to be matched across surveys. The probability of being matched also increases with age, and is higher for nonwhites and for individuals who have at least a high school education. These same results carry over to the probability of having reported earnings, except that the probability is lower for nonwhites and for individuals who own their home.

The results for our baseline wage growth specifications covering all workers in the private non-agricultural sector are given in Appendix Table A3. We present results with and without controlling for sample selection for comparison. Both Mills ratios have positive but imprecisely measured coefficients. In addition, note that exchange rate elasticity is largely unaffected by whether or how we control for sample selection. However, rather than drop both Mills ratios, we continue to control for any possible selection effects arising from having to match individuals across surveys.[13] Finally, the baseline job-changing Probit results are given in Appendix Table A4.

B. Exchange rates and wages. Table 1 provides a summary of the wage elasticities with respect to exchange rates, with workers disaggregated by major industry groups. In the "Full Population" column, the estimated specifications combine all workers (both job-stayers and job-changers) in a particular industry group. We report the average wage elasticities with respect to dollar movements (with depreciations being upward exchange rate movements).[14] Each cell is drawn from a distinct regression specification. For our full sample, a 10 percent dollar depreciation is associated with a 0.8 percent increase in wages (which is consistent with the range of elasticities found in the literature). As suggested by the theory, wages of workers in lower-markup manufacturing industries show more responsiveness to dollar movements than do wages of workers in higher-markup industries.

The next two columns of Table 1 show that there are significant differences in estimated wage elasticities across workers based on their job turnover status. The wages of workers who

household head owns or rents the home. In addition, marital status enters in its level form in the Probits, and in change form in the wage growth specification. The Probit specifications also include year effects which are excluded in the wage growth specification.

[13] Here is a quick summary of the other findings in the baseline specification. Wage growth is declining in age until a worker reaches his mid-thirties, increases until he reaches his early fifties, and then again declines until retirement. College workers experience 3-4 percent faster wage growth than high school dropouts. While married men earn a wage premium, a change in marital status does not generate any immediate effect on wage growth. Finally, wage growth varies positively with aggregate and local demand conditions.

[14] These elasticities are based on specifications that do not allow for time-varying trade orientation of industries. In the section on robustness checks, we allow for changing trade orientation of industries.

stay on their current jobs show little sensitivity to changes in the exchange rate. This evidence suggests that internal labor markets largely insulate workers from external shocks arising from changes in the relative value of the dollar. A key finding – and one that is important for untangling the paradox suggested by prior studies – is that the responsiveness of wages in the full population to dollar movements is driven almost entirely by the impact of these dollar movements on the wages of workers who change jobs. For the full population, a 10 percent dollar depreciation (appreciation) is associated with a 2.2 percent wage gain (loss) for job-changers. The job-changer wage elasticity is larger and more precisely measured for workers in non-manufacturing than for workers in manufacturing.[15]

Table 1 Wage Elasticities with Respect to the Dollar, from Unconstrained Specifications by Industry Group and Job Changing status

Sector Observations (full = stayers + changers)	Full Population	Job Stayers	Job Changers
Private Non-agricultural 114,404 = 90,348 + 24,056	0.083** (0.042)	0.017 (0.057)	0.221** (0.088)
Manufacturing 40,763 = 33,944 + 6,819	0.099 (0.063)	0.031 (0.053)	0.148 (0.208)
High markup 19,205 = 16,333 + 2,872	0.043 (0.098)	–0.013 (0.064)	0.206 (0.310)
Low markup 21,558 = 17,611 + 3,947	0.139* (0.071)	0.061 (0.079)	0.107 (0.225)
Non-manufacturing 73,641 = 56,404 + 17,237	0.074 (0.055)	0.011 (0.067)	0.242** (0.092)

Notes: Reported coefficients are wage elasticities with respect to a dollar depreciation. Each elasticity comes from a distinct regression using only the population of workers associated with that cell. See text for list of control variables. Standard errors are given in parentheses and have been adjusted for any non-independence of observations within a year. ** significant at the 5% level.

C. The Mechanism for Exchange Rate Effects on Wages. The full population wage elasticities with respect to exchange rates reflect both the disparate wage effects for job-stayers and job-changers, as well as the probability that workers change jobs. As shown in equation (8), the full population wage elasticity can be decomposed into the sum of the wage elasticity for job-

[15] This finding is unexpected given the higher trade exposure in manufacturing. We return to this question later

stayers, the probability of job turnover times the impact of exchange rates on the wage consequences of changing jobs, and the wage consequences of changing jobs times the impact of exchange rates on job turnover rates. We turn our attention now to estimating this decomposition.

We present the decomposition of the full population wage elasticities by major industry groups in Table 2. Unlike Table 1 where each elasticity is estimated from a separate regression, in Table 2 each industry group's decomposition (i.e. row of the table) is estimated from a separate regression. We expand our baseline specification to include an indicator for whether a worker changes jobs between the March surveys, and we interact this job-changing indicator with the industry-specific exchange rate. As noted earlier, the elasticity of the probability of changing jobs with respect to the exchange rate is estimated using a selection corrected Probit model.

Table 2 **Decomposition of Channels for Individual Worker Wage Elasticities**

Sample	Implied Overall Elasticity	Job-Stayer Elasticity $\delta\,[\Delta w^{S}]/\delta\Delta rer$	Job-Changers: $P \cdot \delta\,[\Delta w^{C} - \Delta w^{S}]/\delta\Delta rer$	Job-Changers: $[\Delta w^{C} - \Delta w^{S}]\,\delta P/\delta\Delta rer$
Private nonagricultural	0.07	– 0.01 (0.06)	$0.21 \times 0.33^{*} = 0.07$ (0.17)	$-0.23^{**} \times -0.05^{**} = 0.01$ (0.01) (0.02)
Manufacturing	0.08	0.00 (0.06)	$0.17 \times 0.36 = 0.06$ (0.27)	$-0.29^{**} \times -0.08 = 0.02$ (0.02) (0.03)
High Markup	0.04	–0.04 (0.07)	$0.15 \times 0.50 = 0.07$ (0.35)	$-0.27^{**} \times -0.02 = 0.00$ (0.02) (0.50)
Low Markup	0.12	0.03 (0.09)	$0.18 \times 0.28 = 0.05$ (0.27)	$-0.31^{**} \times -0.13 = 0.04$ (0.02) (0.04)
Non-Manufacturing	0.07	–0.01 (0.06)	$0.23 \times 0.30^{**} = 0.07$ (0.14)	$-0.20^{**} \times -0.03 = 0.01$ (0.01) (0.02)

Notes: Δw^{C} denotes the percent change in wages for a job-changer, while Δw^{S} denotes the percent change in wages for a job-stayer. Elasticity decomposition is with respect to a dollar depreciation. Each row is derived from a separate regression, with the implied overall elasticity equal to the sum of the three terms in the decomposition. ** denotes significance at the 5% level. * denotes significance at the 10% level.

when we control for industry trade exposure.

Consider first the decomposition for the private nonagricultural sector as a whole. The average job-changing rate is 21 percent, and the wage growth for changers tends to lag the wage growth of stayers by 23 percent. However, both the incidence and consequence of job-changing are sensitive to shifts in the relative strength of the dollar. A ten percent dollar depreciation lowers the probability of job changing by 0.5 percentage points and reduces the gap in wage growth between changers and stayers by 3.3 percent.

The decomposition shows that job-changing affects the full population wage elasticity primarily through the impact of exchange rates on the wage consequences of job-changing. This basic pattern is repeated in each of the major industry groupings. The implied full population elasticities, which are simply the sum of the three components in the decomposition, match up well with the estimated full population wage elasticities reported in Table 1.

D. Differences By Education Levels of Workers. The results in Table 1 allow for the wage elasticities to vary by broad industry groups, but constrain the elasticity to be equal for all workers with an industry group. We relax this assumption in Table 3 where within each industry group we allow for elasticity differences across three education levels. Specifically, we look at high school dropouts, high school graduates, and workers with at least some college education.

The general pattern that we observed in Table 1, where the full population elasticity is determined primarily by the job-changers, is evident here as well. In addition, we can see that within a broad industry group, the job-changing elasticities tend to be larger for the less skilled (i.e. less educated) workers.[16] For the private nonagricultural sector as a whole, a dollar depreciation is associated with higher wages only for workers with a high school degree or less. Wages of workers with at least some college education are relatively unaffected by exchange rate movements.

In addition, the magnitudes of the elasticities for low-skilled workers who change jobs are quite large. For the private nonagricultural sector as a whole, a ten percent dollar depreciation is associated with a 4.8 percent wage increase for high school dropouts who change jobs, and a 3.7 percent wage increase for high school graduates who change jobs. This basic pattern of elasticities is mirrored across the major industry groups.

[16] Hoynes (1999) documents that less skilled workers have weekly earnings that are more sensitive to general cyclical shocks.

Table 3. Wage Elasticities, by education & industry groups and job changing status			
Sector / Education Attainment Observations (full = stayers + changers)	Full Population	Job Stayers	Job Changers
Private non-agriculture			
High school dropouts 19,539 = 14,166 + 5,373	0.125 (0.080)	–0.087 (0.098)	0.484** (0.122)
High school graduates 46,445 = 36,048 + 10,397	0.156** (0.035)	0.049 (0.061)	0.374** (0.119)
Some college + 48,420 = 40,134 + 8,286	–0.017 (0.038)	0.020 (0.056)	–0.201 (0.118)
Manufacturing			
High school dropouts 8,206 = 6,470 + 1,736	0.256** (0.115)	0.050 (0.075)	0.556 (0.478)
High school graduates 17,738 = 14,665 + 3,073	0.136** (0.059)	0.023 (0.087)	0.253 (0.399)
Some college + 14,819 = 12,809 + 2,010	–0.043 (0.061)	0.021 (0.045)	–0.379** (0.180)
Non-manufacturing			
High school dropouts 11,333= 7,696 + 3,637	0.063 (0.130)	–0.174 (0.133)	0.491** (0.133)
High school graduates 28,707 = 21,383 + 7,324	0.164** (0.042)	0.061 (0.055)	0.412** (0.103)
Some college + 33,601= 27,325 + 6,276	–0.008 (0.049)	0.022 (0.070)	–0.165 (0.151)

Notes: Reported coefficients are wage elasticities with respect to a dollar depreciation. Each elasticity comes from a distinct regression using only the population of workers associated with that cell. See text for list of control variables. Standard errors are given in parentheses and have been adjusted for any non-independence of observations within a year. ** significant at the 5% level.

We expand on this theme in Table 4 where we examine the effect of education on the decomposition of the overall wage elasticity. In general, additional education reduces the likelihood of job turnover, the wage consequences of job turnover, and the impact of exchange rates on both.

Looking at the entire private nonagricultural sector, the incidence of job-changing falls from 27 percent for high school dropouts to 17 percent for workers with at least some college. The wage growth of job changers lags that of job stayers by 30 percent for high school dropouts, but only by 17 percent for workers with some college. A 10 percent dollar depreciation is associated with a 6.7 percent narrowing of this wage growth gap for high school dropouts, but has no impact on this wage growth gap for workers with some college. Similarly, a dollar depreciation reduces job changing rates for high school graduates, but has no impact on job changing rates for more educated workers.

Table 4	Decomposition of Channels for Individual Worker Wage Elasticities By Industry & Education groups			
Sample	Implied Overall Elasticity	Job-Stayer Elasticity	Job-Changers	
			$P \cdot \delta [\Delta w^{C.} - \Delta w^{S}] / \delta \Delta rer$	$[\Delta w^{C.} - \Delta w^{S}]\, \delta P / \delta \Delta rer$
Private nonagricultural				
High school dropouts	0.09	–0.11 (0.10)	0.27 x 0.67** = 0.18 (0.17)	–0.30** x –0.06 = 0.02 (0.01) (0.05)
High school graduates	0.14	0.02 (0.07)	0.22 x 0.45** = 0.10 (0.20)	–0.25** x –0.08** = 0.02 (0.01) (0.03)
Some college +	–0.02	0.00 (0.06)	0.17 x –0.13 = –0.02 (0.19)	–0.17** x –0.01 = 0.002 (0.01) (0.03)
Manufacturing				
High school dropouts	0.23	0.06 (0.07)	0.21 x 0.54 = 0.11 (0.42)	–0.36** x –0.16** = 0.06 (0.02) (0.08)
High school graduates	0.12	–0.02 (0.10)	0.17 x 0.52 = 0.09 (0.54)	–0.31** x –0.15** = 0.05 (0.02) (0.05)
Some college +	–0.04	–0.01 (0.05)	0.14 x –0.20 = –0.03 (0.24)	–0.21** x 0.05 = –0.01 (0.01) (0.05)
Non-Manufacturing				
High school dropouts	0.03	–0.21 (0.13)	0.32 x 0.74** = 0.24 (0.19)	–0.26** x 0.004 = –0.001 (0.02) (0.07)
High school graduates	0.15	0.04 (0.05)	0.26 x 0.40** = 0.10 (0.14)	–0.22** x –0.04 = 0.01 (0.01) (0.04)
Some college +	–0.005	0.01 (0.07)	0.19 x –0.11 = –0.02 (0.21)	–0.16** x –0.03 = 0.005 (0.02) (0.05)

Notes: Δw^C denotes the percent change in wages for a job-changer, while Δw^S denotes the percent change in wages for a job-stayer. Elasticity decomposition is with respect to a dollar depreciation. Each row is derived from a separate regression, with the implied overall elasticity equal to the sum of the three terms in the decomposition. ** denotes significance at the 5% level. * denotes significance at the 10% level.

These results demonstrate that there is considerable diversity across skill groups in the impact of a change in the relative strength of the dollar. The full population elasticity for a major industry group is a poor proxy for the elasticity for low skilled workers within that industry. In addition, the evidence shows that even within a specific skill group, those workers who are induced to change jobs typically shoulder the burden imposed by a strong dollar. There is little evidence that workers who stay with their same employer are significantly impacted by a strong dollar.

E. Robustness checks. In this section we consider interacting the industry exchange rates with alternative trade exposure measures. Trade exposure has steadily increased over our sample period as the U.S. economy has become more open. For example, consider the trade exposure measure defined as the sum of the industry export and import trade shares. This trade exposure measure increased from an average of 6.6 percent in 1977 to 10.6 percent in 1997 (weighted by the number of workers in each industry). The maximum industry value increased from 23.8 percent in 1977 to 82.8 percent in 1997. There is also important variation in trade exposure across industries. While the overall average trade exposure in our sample is 8.4 percent, the average is 20.3 percent for manufacturing industries and 1.9 percent for non-manufacturing industries.

Our results presented so far show that wage elasticities with respect to dollar fluctuations vary by industry structure and worker skill levels. The data also show that these wage elasticities vary with industry trade exposure, here defined as the time-varying sum of industry-specific export and import shares. Appendix Table A5 summarizes the variation in implied wage elasticities for different percentiles of trade exposure within the specific industry group. For our entire sample the inner quartile range of wage elasticities is 0.11, while the difference between the 90th and 10th percentiles is 0.26.[17]

A puzzle raised in Table 1 is the average non-manufacturing wage elasticity of 0.07. This would seem too high given the low average trade exposure for non-manufacturing. What we learn from Appendix Table A5 is that the median wage elasticity in non-manufacturing is only 0.01, significantly below the average of 0.07. In contrast, the 90th percentile wage elasticity is 0.151, even though the trade exposure for this percentile is only 6.4 percent. The data suggest that the sensitivity of non-manufacturing wages to dollar fluctuations rises rapidly with the

[17] Measuring trade exposure by summing the export and import shares is relevant if the primary effects of exchange rate movements are on producer revenues derived from both foreign sales and domestic sales due to import penetration of domestic markets. An alternative measure is the *difference* between the industry export and import shares. Empirically the import penetration in an industry and the industry's use of imported inputs are highly correlated and similarly scaled. See Campa and Goldberg 1997. When we redefine trade exposure to be the difference between the industry export and import trade shares, we find that the precision of the interaction term with the industry exchange rate generally declines. In addition, the implied sensitivity of the wage elasticity to variations in this measure of trade exposure is higher.

measured trade exposure. Each percentage point increase in trade exposure increases the non-manufacturing wage elasticity by 0.2, while a similar increase in trade exposure increases the manufacturing wage elasticity by only 0.05. What we can not tell from our data is whether this higher trade sensitivity of non-manufacturing wages is real, or whether it reflects underlying measurement problems with the non-manufacturing trade share data.

V. Summary and Conclusions

We explore in this paper the magnitude of wage sensitivity to dollar movements and the channels for this sensitivity, through a combination of adjustments to wages by: (1) workers who remain with their same jobs, (2) workers who receive an alternative wages after switching jobs, or (3) induced changes in the frequencies of switching by particular classes of workers in a specific industry groups.

We have demonstrated that the process of aggregation masks large and significant wage responses to dollar movements within particular industry groups. When individuals are differentiated by their educational attainment, we find that workers without high school diplomas always have more sensitive wages than more educated workers. Furthermore, the effects of dollar movements are mainly shouldered by job-changers. Those workers who change jobs suffer larger adjustment penalties as opposed to workers who remain on their jobs. While dollar movements do have some effect on job-turnover rates, the size and importance of this channel is not substantial for the average workers in specific industry groups.

Overall, this paper has demonstrated that dollar fluctuations translate into significant wage effects for specific groups of workers in the U.S. economy. Many of these distributional effects are lost in studies of aggregate data, either because of changes in the composition of the workforce or because of aggregation problems. The results from aggregate studies can give misleading conclusions when applied to the question of the actual disruptiveness of dollar fluctuations for different skill classes of workers.

We also provide a possible answer to the paradox presented by the aggregate studies that show larger wage than employment elasticities. This result seems inconsistent with the perception that relatively more of the adjustment to dollar fluctuations likely occur on the employment rather than the wage margin. Our findings show that for workers who remain with their same employer, there is little significant wage response to dollar movements. However, there are large wage responses for workers who change employers. Dollar movements affect both the incidence and wage consequences of job turnover, with the magnitudes of these effects increasing as one moves down the skill spectrum.

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Appendix

Computation of Permanent Exchange Rates: The Beveridge-Nelson procedure decomposes an I(1) time series into its transitory and permanent components. We follow Beveridge and Nelson (1981) and Huizinga (1987), and model the exchange rate decomposition using lags of the real exchange rate. In particular, we assume that the first differences of the quarterly (log) real exchange rate follow an AR(4) process, so that the transitory departure of the real exchange rate from its expected long-run equilibrium, e_t^{BN}, is given by:

$$E_t\left(e_t^{BN}\right) = -E_t\left(\sum_{j=1}^{\infty} \Delta e_{t+j} \mid \Delta e_t, \Delta e_{t-1}, \Delta e_{t-2}, \Delta e_{t-3}\right)$$

The actual variance decomposition results suggest that the temporary component of exchange rate changes accounts for only a small proportion of the variance of the real exchange rate series.

Appendix Table A1. Description of the Population of Male Workers

	All Observations Variable Means (percent of population)		Job Changers Only Variable Means (percent of population)	
	Unmatched	Matched	Unmatched	Matched
Manufacturing	.32	.36	.25	.28
Higher Markup Industries	.15	.17	.10	.12
Lower Markup Industries	.17	.21	.15	.16
Non-Manufacturing	.68	.64	.75	.72
Job Changers	.30	.19		
High School Dropouts	.20	.17	.25	.23
High School Graduates	.40	.41	.41	.43
Some College	.20	.20	.19	.19
College Graduates	.20	.22	.15	.15
Age in years (standard deviation)	36.9 (11.9)	38.6 (11.4)	33.1 (12.1)	36.1 (12.2)
Race (Non-white)	.10	.12	.11	.13
Married	.69	.76	.55	.67
Own home	.67	.78	.55	.71
Number of Observations	204,690	114,404	62,503	24,056

Notes: **All statistics for Matched CPS use base year values. Unmatched sample excludes individuals with missing earnings.

Appendix Table A2. Probability of Matching Individual Across Surveys & Nonmissing Wage		
	Match Across Surveys	Nonmissing Wage
Variable	Probit Coefficient (Standard Error)	Probit Coefficient (Standard Error)
Married	0.179** (0.004)	0.366** (0.007)
Unmarried kids < 18 years old	0.049** (0.007)	0.048** (0.007)
Unmarried kids < 6 years old	–0.020** (0.008)	0.107** (0.009)
Age of household head	0.998** (0.028)	0.509** (0.029)
Age squared	–0.040** (0.001)	–0.020** (0.001)
Age cubed (x100)	0.071** (0.002)	0.035** (0.002)
Age fourth (x1,000)	–0.004** (0.000)	–0.002** (0.000)
High school graduate	0.148** (0.007)	0.201** (0.007)
Some College	0.093** (0.008)	0.260** (0.008)
College Graduate	0.108** (0.008)	0.260** (0.008)
Nonwhite	0.018** (0.008)	–0.306** (0.008)
Own a home	0.619** (0.005)	–0.113** (0.006)
Year effects included	Yes	Yes
Number of observations	286,471	286,471
Notes: Sample restricted to individuals who are "at risk" of being matched in the subsequent March CPS survey (includes individuals with missing earnings). Probit coefficients are given with standard errors are given in parentheses. ** denotes significant at the 5% level		

Appendix Table A3.	**Wage Growth – Baseline Specification**		
Variable	(1)	(2)	(3)
Age	–0.113** (0.003)	–0.115** (0.018)	–0.114** (0.011)
Age squared (x10)	0.027** (0.001)	0.028** (0.005)	0.028** (0.003)
Age cubed (x1,000)	–0.021** (0.001)	–0.022** (0.004)	–0.022** (0.002)
Becomes married	0.005 (0.008)	0.006 (0.009)	0.006 (0.009)
Becomes single	–0.006 (0.009)	0.014 (0.014)	0.002 (0.009)
High school graduate	–0.002 (0.005)	0.010 (0.006)	0.004 (0.005)
Some college	0.008 (0.008)	0.019** (0.008)	0.012 (0.008)
College graduate	0.029** (0.007)	0.041** (0.007)	0.034** (0.007)
% change in real GDP	1.197** (0.121)	1.157** (0.104)	1.179** (0.119)
Change in local relative demand conditions	0.585** (0.168)	0.586** (0.165)	0.585** (0.166)
% change in real exchange rate	0.085** (0.032)	0.087** (0.029)	0.083** (0.033)
Mills ratio – Match across surveys		0.052 (0.058)	0.048 (0.054)
Mills ratio – Nonmissing wage		0.059 (0.091)	
2-Digit industry effects included	Yes	Yes	Yes
N = 114,404			

Notes: Standard errors are given in parentheses. ** denotes significant at the 5% level. * denotes significant at the 10% level.

Appendix Table A4. Job Changing Probit	
Variable	Probit Coefficient
Prior job changing	0.957** (0.009)
Age	0.013 (0.013)
Age squared (x100)	-0.111** (0.034)
Age cubed (x10,000)	0.141** (0.028)
Becomes married	-0.127** (0.026)
Becomes single	0.048* (0.026)
High school graduate	-0.123** (0.012)
Some college	-0.167** (0.013)
College graduate	-0.301** (0.014)
% change in real GDP	-0.893** (0.192)
Change in local relative demand conditions	-0.778** (0.251)
% change in real exchange rate	-0.201** (0.076)
N=114,404	

Notes: Probit coefficients with standard errors given in parentheses. ** denotes significant at the 5% level. * denotes significant at the 10% level.

Appendix Table A5.	**Wage Elasticities: by Trade Exposure Levels**				
	Percentile of trade exposure				
Sector	10^{th}	25^{th}	50^{th}	75^{th}	90^{th}
Private nonagricultural					
Trade exposure	0.002	0.003	0.023	0.119	0.265
Wage elasticity	0.001^{**} (0.0005)	0.002^{**} (0.0005)	0.013^{**} (0.005)	0.066^{**} (0.027)	0.147^{**} (0.061)
Manufacturing					
Trade exposure	0.060	0.095	0.169	0.268	0.412
Wage elasticity	0.029^{*} (0.015)	0.046^{*} (0.024)	0.081^{*} (0.043)	0.129^{*} (0.069)	0.199^{*} (0.106)
Non-manufacturing					
Trade exposure	0.002	0.003	0.004	0.021	0.064
Wage elasticity	0.005^{**} (0.002)	0.006^{**} (0.003)	0.010^{**} (0.005)	0.049^{**} (0.023)	0.151^{**} (0.071)

Notes: The trade exposure for a given sector and percentile represents the worker weighted level of trade exposure. The wage elasticity reported for a given percentile is estimated at the indicated trade exposure level for that sector. Standard errors are given in parentheses. ** denotes significance at the 5% level. * denotes significance at the 10% level.